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W-133

Benefits and Costs of Resource Policies Affecting Public and Private Land

Eleventh Interim Report

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INTRODUCTION

This volume contains the proceedings of the 1998 W-133 Western Regional Research Project Technical Meeting on "Benefits and Costs of Resource Policies Affecting Public and Private Land". The meeting took place March 18-20, in Colorado Springs, Colorado. Despite complications introduced by 18 inches of snow on the opening day of the conference, over 50 scientists managed to arrive and participate in the meeting.

The current version of the W-133 Regional Research Project officially began on October 1, 1997, but most of the project scientists have worked together on previous regional projects. The goals of the current project arose from the questions raised by past research and from recognition of newly emerging problems. The project objectives are to address problems associated with:

1. Valuing Changes in Recreational Access
2. Benefits and Costs of Agro-environmental Policies
3. Benefits Transfer for Groundwater Quality Programs
4. Valuing Ecosystem Management of Forests and Watershed

Nineteen Experiment Stations are official project participants, while other stations sent a representative to the technical meeting. The W-133 Technical Meeting also benefitted from the expertise and participation of scientists from the United States Departments of Agriculture, Interior, and Commerce, as well as the Environmental Protection Agency. Further, a number of private consulting firms involved in natural resource damage assessment also attended and participated in the technical meeting.

The current volume is organized around the goals and objectives of the project, although many of the studies could overlap more than one goal. For example, a number of papers introduce innovative theoretical approaches or methodological techniques which could easily apply to more than one project objective. As such, classification of a paper under a particular objective might not reflect the full contribution of the research it reports. However one might wish to arrange them, the papers contained in this volume demonstrate the rich variety of high quality research produced by the cooperating scientists of the W-133 Regional Research Project.

Finally, I would like to thank Mrs. Debbie Sharp for assistance in preparing this volume.

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Rain soaks East, snow clobbers West

The Associated Press

Showers and thunderstorms hit most of the nation east of the Mississippi River on Wednesday, while snow and strong winds plagued the Rocky Mountains.

Some severe storms produced large hail in parts of Mississippi and Louisiana, but no significant damage was reported. Light to moderate rain fell farther north. Freezing rain early in the morning in Wisconsin turned to rain or snow later in the day.

A storm system located across Colorado tracked eastward and strengthened throughout the day. Snow combined with strong winds to produce near-blizzard conditions in Wyoming, Colorado and northern New Mexico.

Windblown snow, falling at an average of an inch an hour in some spots of Colorado, reduced visibility.

Woody Wright of Silver City Towing in Idaho Springs said his company went from call to call to help stranded travelers.

"It's real slick, wet and heavy, and a lot of it," Wright said.

"This ranks right up there with that blizzard we had in October," said Colorado State Patrol spokesman Scott Nathlich. "The snow is falling harder than anybody thought. The snowplows can't keep up."

By Wednesday afternoon, up to 14 inches of snow had hit the west side of Denver, according to Paul Gard of the National Weather Service.

"March is normally our snowiest month. But we were prepared for more snow because of El Nino," he said.

The Colorado Springs Municipal Airport, about 70 miles south of Denver, closed, and there were some delays at Denver International Airport.

Sixty snowplows were sent to clear roads in Denver and another 40 were working on Interstate 70, west of Denver.

In Wyoming, a 110-mile stretch of Interstate 80 between Rawlins and Rock Springs was among the many highways closed. The roads that remained open were described as treacherous.

"Things are just not moving real fast because people are just socked in," said Tierney King-Franz, spokeswoman for the Fremont County Sheriff's Department.

Light to moderate rain showers also associated with the system fell across New Mexico and Arizona. Other parts of the West experienced fair and dry weather.

Heavy rains in Oklahoma destroyed several homes that crumpled into the Cimarron River.

▶ Full weather, 16A

Towards Globally Flexible Estimation of Recreation Demand

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Prepared for the USDA Regional Research Project W-133 Technical Meeting
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Abstract: Towards Globally Flexible Estimation of Recreation Demand

This research seeks to produce more robust, less biased estimates from the two-constraint recreation demand model by applying globally flexible estimation techniques to travel cost data for California whale watching trips. While locally flexible functional forms can improve upon more commonly used forms for estimation of travel cost models, specification errors occur unless the chosen form happens to coincide with the true unknown underlying form. A globally flexible functional form, however, can consistently approximate the true function and its derivatives for all points in the sample range. In this paper, we use a globally flexible functional form to improve accuracy of estimates from the two-constraint recreation demand model.

The empirical model is based on a construction by Chalfant (1987), which combines Deaton and Muellbauer's AIDS model with the Fourier flexible form of Gallant. The resulting functional form preserves the aggregation properties of the PIGLOG class of preferences while approximating the true function to an arbitrary degree of precision. A comparison of model estimation results shows that the locally flexible AIDS model results in specification error.

Introduction

Traditional demand analysis tools do not necessarily generalize to the case of recreation demand due to the lack of observable market and price signals associated with its consumption. One important aspect of recreation is the time involved in a recreation trip. Since time can influence consumption decisions concerning whether to undertake the activity, and also the level of participation in the activity, it creates an additional choice constraint for the consumer. First suggested by Hotelling, and formalized by Clawson in 1959, was the notion that travel time and costs could be used to estimate recreation demand models. This method, based on the notions of revealed preference, became formalized as the travel cost method (TCM), which, in its many variations, is widely used for recreation demand modeling and estimation.

A significant development in recreation demand modeling is the two-constraint framework in which choices are made under both a money and time constraint (Bockstael *et al*; McConnell; Larson; Larson and Shaikh; Smith). However, since empirical specification of these models is complex and data sets can be limited in scope, assumptions are often imposed to simplify estimation. This can result in empirical models which lack a consistent link to the theoretical model defining the decision-making process of the consumer. It has recently been shown that the presence of the time constraint in the two-constraint recreation demand model provides additional structure to relate time and money variables (Larson and Shaikh). These relationships can be used to improve empirical specification of the two-constraint model.

It is possible to maintain notions of consumer demand and preference theory through the theory of the two-constraint model. An empirical model, based on the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer, is constructed using restrictions posed by the theoretical model. We incorporate a Fourier expenditure function, which uses global approximations to improve consistency, to enhance the flexibility of the empirical model. The benefits of global flexibility over local flexibility are examined in the recreation demand case.

The Two-Constraint Model

The omission of time in recreation demand models as a relevant cost in the decision-making process leads to incomplete prices and budgets. Cesario and Knetsch indicated that to avoid correlation between travel cost and travel time variables, it is advantageous to create "full" costs and "full" budgets by combining all travel costs and time costs. These theoretical concepts from the two-constraint utility maximization model can help identify the appropriate specification of the corresponding empirical demand functions.

The two-constraint model uses theoretical developments Smith (1986) and Bockstael *et al.* (1987). The indirect utility function is:

(1) $$V(\mathbf{p},\mathbf{t},z,M,T) = \max_{\mathbf{x}} \; u(\mathbf{x},z) + \lambda(M - \mathbf{px}) + \mu(T - \mathbf{tx})$$

where *u(x,z)* is a twice continuously differentiable utility function, **x** is a vector of consumption goods with money prices **p**, and time prices **t**. Choices are made subject to a money budget constraint M = **px** and a leisure time budget constraint T = **tx**, where T=(total time - work time). Assume that both constraints bind, which implies non-satiation, and that all time must be spent in some activity. One interpretation that would allow for binding constraints relies on the presence of slack variables, where one component of the **x** vector would represent savings, and another component would represent leisure time uncommitted to any activity. Utility is also a function of the quality variable z.

The Lagrange multipliers λ and μ, represent the marginal utility of income and the marginal utility of leisure time, respectively. Therefore, the money value (or, opportunity cost in dollars) of time is defined as the ratio of Lagrange multipliers, $\rho = \mu/\lambda = V_T/V_M$, where V_x, denotes the partial derivative of (1) with respect to the subscripted variable (x in this case).

The two constraints imply two dual expenditure functions. The money expenditure function, $E^M(\mathbf{p},\mathbf{t},z,T,u)$, and the time expenditure function, $E^T(\mathbf{p},\mathbf{t},z,M,u)$, are defined as:

(2) $$E^M(\mathbf{p},\mathbf{t},z,T,u) \equiv \min_{\mathbf{x}} \mathbf{px} \ \textit{subject to} \ T = \mathbf{tx} \ \textit{and} \ u = u(\mathbf{x},z), \text{ and}$$

(3) $$E^T(\mathbf{p},\mathbf{t},z,M,u) \equiv \min_{\mathbf{x}} \mathbf{tx} \ \textit{subject to} \ M = \mathbf{px} \ \textit{and} \ u = u(\mathbf{x},z),$$

respectively. From the dual expenditure functions, the marginal value of leisure time is defined as $\rho = -\partial E^{\mathbf{M}}/\partial T = -\left[\partial E^{\mathbf{T}}/\partial M\right]^{-1}$. This is interpreted as the change in money expenditure needed from a change in the leisure time budget to keep utility constant, or the inverse of the change in the leisure time budget needed from a change in money income to keep utility constant.

From the comparative statics of the model, the Marshallian demands, can be recovered through two separate versions of Roy's Identity:

$$x_i(\mathbf{p},\mathbf{t},z,M,T) \equiv -V_{p_i}/V_M \equiv -V_{t_i}/V_T$$

Larson and Shaikh (1997) show that from these identities, relationships exist between demand coefficients for money price and time price, as well as, for money budgets and time budgets. These relationships will hold regardless of whether or not he individual is trading at the margin or has a fixed work week. In either case, they show that a sufficient condition uses the marginal value of time ρ, to express the demand arguments as full prices, $p_i = (p_i + \rho * t_i)$ and full budgets, $(M + \rho * T)$. This logic is carried over to the empirical model to maintain consistency with the theoretical developments. By constructing full prices and budgets, the dual expenditure functions (2) and (3) are reduced to one "full" expenditure function. Hereafter, all prices and budgets and the expenditure function are defined as "full".

Locally Flexible Forms

Model implications can be ambiguous if restrictive functional forms are used. Therefore, it is beneficial to choose an integrable and flexible functional form. An integrable form uses imposed demand theory restrictions to ensure that the indirect utility function is recoverable from the demand equations. A flexible form contains enough model parameters to provide a local approximation at particular values of the price-income ratios (Deaton and Muellbauer). While locally flexible forms have been used extensively for demand analysis, only recently have they been used to estimate recreation demand (e.g. Creel; Larson, *et al.*).

The Almost Ideal Demand System

One of the advantages of the AIDS model over more commonly used empirical models of recreation, is its aggregation properties, which are consistent with the price-independent, generalized-logarithmic (PIGLOG) class of preferences. PIGLOG preferences permit exact aggregation over consumers so demand can be analyzed for a rational, representative consumer. Deaton and Muellbauer define the PIGLOG cost function as:

(4) $$\log c(u,\mathbf{p}) = (1-u)\log\{a(p)\} + u\log\{b(p)\}.$$

The AIDS model is defined as flexible since it contains enough parameters so that at a single point, the derivatives are equal to those of an arbitrary cost function. It is defined as:

(5) $$\log a(p) = a_0 + \sum_k a_k \log p_k + \tfrac{1}{2}\sum_k \sum_j \gamma^*_{jk} \log p_k \log p_j \text{, and}$$

$$\log b(p) = \log a(p) + \beta_0 \prod_k p_k^{\beta_k}.$$

The resulting Marshallian share is:

(6) $$s_i = \alpha_i + \sum_j \gamma_{ij} \ln p_j + \beta_i \ln(M/P),$$

where log P = log a(p) and M is total expenditure. Elasticities in the linear approximate AIDS model are calculated as $\hat{\varepsilon}_{ii} = -1 + \hat{\gamma}_{ii}/\hat{s}_i - \hat{\beta}_i$, for own price, $\hat{\eta}_i = 1 + \hat{\beta}/\hat{s}_i$, for income, and $\hat{\varepsilon}_{iz} = \hat{\gamma}_{iz}/\hat{s}_i$, as the elasticity with respect to quality.

Limitations of Local Flexibility

While the AIDS model provides both flexibility and integrability, we have no information about the specific form of utility function for individual or aggregate consumption. White (1980) argued that the estimates of a local approximation generally do not correspond to the true function or its derivatives at any point in its domain, unless the parametric specification happens to coincide with the true unknown underlying function. Nevertheless, even with potential limitations, the AIDS model has many beneficial properties for consumer demand analysis and provides a simple framework for estimation of share equations.

Global Flexibility and the Fourier Flexible Form

Gallant (1981) defines a measure to be globally flexible if it satisfies the requirements of the *Sobolev norm.* Following Elbadawi, *et al.* (1983), as a compact notation for partial differentiation, let

$$D^{\psi}f(x) = \frac{\partial^{\psi}}{\partial x_1^{\psi}\partial x_2^{\psi_2}\ldots\partial x_k^{\psi_k}} f(x)$$

where the function f(x) is a K x 1 vector, ψ is a K x 1 vector of nonnegative integers and $|\psi| = \sum_{i=1}^{k}|\psi_i|$, is the order of the partial derivative, so that when ψ is the zero vector, $D^{\psi}f(x) = f(x)$. The *Sobolev norm* of the function f(x) is defined as,

$$\|f(x)\|_{m,X} = \max_{|\psi|^{*}\leq m} \sup_{x\in X}|D^{\psi}f(x)|,$$

where *m* is a nonnegative finite integer and X is the domain of the function f(x). Sobolev flexibility then, is global, since the supremum is over X. The Sobolev norm of the difference between the true function and the approximating model $H(x,\phi)$, (where x is a vector of exogenous variables and ϕ is the parameter vector) tends almost surely to zero:

$$\|f^{*}(x) - H(x,\hat{\phi})\|_{m,X} \rightarrow 0, \text{ almost surely.}$$

A Fourier series approximation, a linear combination of sine and cosine functions, possesses Sobolev-flexibility. Because it has the ability to exactly represent any true function, a flexible functional form using a Fourier series should avoid the restrictive nature and limitations of a locally flexible functional form (Chalfant). For exact representation of a function, a Fourier series may require an infinite number of expansion terms. However, since data sets are finite, exactness is not generally feasible, so that it is appropriate to use a Fourier series to approximate a function to an arbitrary degree of precision.

The Fourier form of Gallant (1981, 1982, 1984), uses a leading quadratic term with a multivariate Fourier series expansion. Generally, it can be written as:

$$(7) \qquad H(x|\phi) = u_0 + \mathbf{b}'\mathbf{x} + \tfrac{1}{2}\mathbf{x}'\mathbf{C}\mathbf{x} + \sum_{\alpha=1}^{A}\{u_{0\alpha} + 2[u_\alpha \cos(\lambda \mathbf{k}_\alpha' \mathbf{x}) + v_\alpha \sin(\lambda \mathbf{k}_\alpha' \mathbf{x})],$$

where the $\mathbf{k}_\alpha$'s are N x 1 vectors of positive and negative integers, λ is a scaling factor and $\mathbf{x}$ is a vector of logged prices and income. C is an N x N matrix defined by $C = -\sum_{\alpha=1}^{A}\lambda^2 u_{o\alpha} k_\alpha k_\alpha'$ and the vector of parameters is $\phi = (u_0, \mathbf{b}, u_{01}, u_1, v_1, \ldots, u_{0A}, u_A, v_A)$. Using the second order polynomial allows us to test if the Fourier form reduces to a more simple, locally flexible form, such as the AIDS.

The Fourier flexible form has been used to improve specification bias on elasticities (Chalfant and Gallant) and welfare estimates (Creel; Chen and Randall; Creel and Loomis). Chalfant (1987) combined a Fourier cost function with an AIDS specification resulting in a globally flexible functional form, that included the aggregation properties of the AIDS model. Following Chalfant's model, we construct a globally flexible model which is specified in the context of the recreation model by using full prices and budgets as defined previously, and by treating the time conversion factor (based on the reported wage) as an endogenous parameter to be estimated.

The Empirical Model

The empirical model replaces log a(P) from equation (5), of the AIDS model with the Fourier cost function in (7). The resulting "full" expenditure function is:

$$\log e(p,z,u) = u_0 + \mathbf{b}'\mathbf{x} + \tfrac{1}{2}\mathbf{x}'\mathbf{C}\mathbf{x} + \sum_{\alpha=1}^{A}\left\{u_{0\alpha} + 2[u_\alpha \cos(\lambda \mathbf{k}_\alpha{}' \mathbf{x}) + v_\alpha \sin(\lambda \mathbf{k}_\alpha{}' \mathbf{x})]\right. + u\beta_0 \prod_k p_k^{\beta_k} \tag{8}$$

The Marshallian share equation is of the form:

$$s_i = b_i - \lambda \sum_{\alpha=1}^{A} \{u_{0\alpha} \lambda \mathbf{k}' \mathbf{x} + 2[u_\alpha \sin(\lambda \mathbf{k}' \mathbf{x}) + v_\alpha \cos(\lambda \mathbf{k}' \mathbf{x})]\}\mathbf{k} \;\; + B_1 \ln(M/P). \tag{9}$$

Own price elasticity is defined as $\hat{\varepsilon}_{ii}^{GF} = (-\lambda^2 \hat{(}\gamma_{ii} + \hat{u}_p \cos(\lambda \ln p_i) - \hat{v}_p \sin(\lambda \ln p_i))/\hat{s}_i - 1$, and elasticity with respect to quality as $\hat{\varepsilon}_{iz}^{GF} = (-\lambda^2 (\hat{\gamma}_{1z} + \hat{u}_z \cos(\lambda \ln z) - \hat{v}_p \sin(\lambda \ln z)) / \hat{s}_i$. Income elasticity is calculated as in locally flexible model since the Fourier cost function does not affect the income term. The combined, globally flexible model reduces to the locally flexible AIDS model if the $u_{0\alpha}$'s correspond to the γ_{ij}'s from the AIDS and the parameters on the expansion terms, the u_α's and v_α's, are all equal to zero. By estimating the γ_{ij}'s directly in the unrestricted share equation in (7), a simple F test can be used to test reduction of the globally flexible model to the locally flexible one (Piggott 1997).

California Whale Watching

The data used to illustrate the model are from on-site intercepts of whale-watchers at four sites in California during the winter of 1991-92. The survey collected information on trips taken so far that season, expected future trips, travel time, travel costs, and whether or not the trip was their primary destination. It also collected data on actual contributions to marine mammal groups, time spent reading, watching, or thinking about wildlife and whales, as well as purchases of whale-related merchandise. Respondents were asked to rate the relative importance of viewing whales, knowing whales exist, existence of whales for future generations, etc. Demographic information including work status and job structure related variables, wage rates, and income was also collected

Four goods can be used to define the time and money share systems from the whale-watching data set: whale-watching trips (x_1); time donations to whale- and marine mammal-related organizations (x_2); time volunteered for such organizations (x_3); and a numeraire good for consumption of all other goods (x_4). Each good has an associated money and time price which are used to construct a full price with the time conversion factor to be estimated. In some cases, the money or time price of a good may be zero or one. The time price of a money donation is negligible and assumed to be zero. The time price of a time donation is taken to be one, although realistically it could be greater than one due to transaction costs from driving to the site, etc. The money cost of a time donation represents transaction costs of donating time. The survey does not provide information on this variable but it is suspected to be small, thus it is taken to be zero.

In addition to the time and money prices, it is expected that the individual's whale-watching success will influence both trips demand and, potentially, the willingness to make donations of time and money. The quality variable (z) is the individual's ex-ante expectation of whale sightings for the whale-watching trip when they were contacted. Money budget (M) is the household income before taxes, and the time budget is amount of non-working time in the number of weekend and paid vacation days.

Model Estimation

Using SAS Version 6.12, incomplete demand systems using the trips (s1) share equations from (6) and (9) were estimated as the restricted and unrestricted models, respectively. Intercept dummies were included to account for differences in whale watching sites. The x vector is comprised of logged full own price (lp1f), logged full prices for time donations (lp2f) and money donations (lp3f), and logged quality (lz). The k_α terms are defined such that only first order Fourier expansions on own price and quality are used.

McConnell and Strand (1981) defined the opportunity cost of time as a constant times average income. We construct full prices and budgets similarly by defining the marginal value of time as, $\rho = v * w_i$, where v is a parameter to be estimated, and w_i, is the individual's reported wage. The value of v is not restricted to be less than one since the wage rate is not necessarily the upper bound of the value of leisure time (Bockstael, *et al.;* Larson; Larson *et al.*).

Preliminary Results

Table 1 provides the estimated parameters and standard errors for both models. The value of v, which when multiplied by the reported wage, converts time to money units, is significant in both models. The value of v is .65 in the restricted model and .9 in the unrestricted, indicating that the wage is greater than the implied value of time. This finding is consistent with other empirical studies, which have treated wage as the upper bound of the value of time (Smith, *et al.*; McConnell and Strand). The value of time varies by individual wage rates.

Both models found own price significant, however, none of the cross price estimates (time and money donations) are significant in either model and are omitted from the reported results for brevity. The first site dummy variable in the locally flexible model is significant, indicating possible variation by site choice. Two of the Fourier expansion terms are significant in the globally flexible model. While the coefficient estimates on the expansion terms are not interpretable on their own, the significance suggests their necessity in correctly approximating the true function. Quality (z) is also a significant in the globally flexible model.

Table 2 provides elasticities at the mean and median predicted shares, as well as the range of values for both models. The signs and magnitudes are as expected given the nature of the data. There does appear to be more variation in elasticities in the globally flexible model, although it is difficult to compare since income and quality effects were not significant in the locally flexible model. Whale-watching trips are own-price inelastic in both models, but close to one in the locally flexible model. Trips appear to be income inelastic, and normal, although roughly 8% of calculated elasticities are negative and the coefficient estimate borders on insignificant. The elasticity of trips with respect to quality is positive and less than one implying an inelastic effect.

Testing the AIDS Specification

A main objective of this research is to test for possible misspecification from choosing a fully parametric specification. By setting the Fourier expansion terms coefficients v_z, u_z, v_p, and u_p, to zero, the globally flexible model reduces to the locally flexible AIDS model. The null hypothesis that the Fourier expansion terms are zero is rejected ($F_{4,277} = 19.945$). This implies that under the PIGLOG preference assumption, the correct specification of the empirical model requires the Fourier expansion terms. In this case, the AIDS model does not correctly approximate the true underlying model or its derivatives, thereby resulting in specification error and biased estimates

Summary and Further Research

This research improves upon traditionally used estimation methods for recreation demand by using a flexible and integrable functional form to link the theoretical and empirical models. However, we do not know how well the chosen form approximates the true underlying function. The AIDS specification of the model is tested through estimation of a globally flexible model, which approximates the true unknown function to an arbitrary degree of precision. Under the maintained hypothesis of PIGLOG preferences, preliminary results show that the AIDS model results in specification error, implying that the chosen locally flexible model does not coincide with the true function.

The value of time to varies with wage rate. However, it was treated as a constant fraction across the sample. Ultimately, the opportunity cost of time varies by individual. Since demand analysis generally involves the "representative consumer", defining this empirically has proven quite problematic and requires future research.

Several improvements to the estimation are necessary to improve model consistency and reliability of estimates. Both models were estimated using Stone's price index, which is a useful simplification but can lead to biased elasticity estimates (Buse 1994). Also, the globally flexible model is estimated using only a first order expansion. Higher order effects should be explored by including additional Fourier expansion terms. Gallant has likened the choice of order to choosing the lag in a time series model (1982). By including additional expansion terms, simple Wald tests can be used to determine the optimal order.

In conclusion, this research has explored the differences in global and local approximations to an unknown underlying function. The preliminary results from this research warrant caution in interpreting elasticity or welfare estimates resulting from parametric models.

Table 1: Estimation Results

(291 observations, standard errors in parentheses)

Variable	*Coefficient*	*Locally Flexible Estimate*	*Globally Flexible Estimate*
Intercept	α_1	-0.0003	0.002389
		(.000711)	(.001694)
Wage Fraction	K	0.659304***	0.90602**
		(.20679)	(.39173)
Site 1 Dummy	D1	4.00E-06**	9.53E-06
		(1.95E-06)	(2.89E-05)
Site 2 Dummy	D2	-1.49E-07	-2.8E-04***
		(3.47E-06)	(.000105)
Site 3 Dummy	D3	-4.72E-06	-2.9E-04***
		(4.62E-06)	(8.49E-05)
Own Price	γ_{11}	1.07E-05***	-1.04E-03***
		(1.49E-06)	(.000194)
Quality	γ_{1Z}	3.32E-07	-.12E-04**
		(3.42E-07)	(5.06E-05)
Income	$\beta 1$	4.13E-05	-2.82E-04
		(7.96E-05)	(.000184)
Quality Sine Expansion	u_z		5.98227E-06
			(6.89E-05)
Quality Cos Expansion	v_z		9.16E-05**
			(4.06E-05)
Own Price Sine Expansion	u_p		1.575E-05
			(8.8E-05)
Own Price Cos Expansion	v_p		-4.3E-04***
			(.000115)

*significant at 10%, **significant at 5%, ***significant at 1%.

Table 2: Calculated Elasticities

	Locally Flexible Model				*Globally Flexible Model*			
Elasticity	At Med. Share	At Mean Share	Min	Max	At Med. Share	At Mean Share	Min	Max
ε_{11}	-0.97827	-0.98127	-0.99418	-0.40231	-.53821	-.646358	-.89015	1.924619
ε_{1z}	0.000663	0.00058	0.000181	0.018481	.104688	.078308	.019401	.475676
η_1	1.083702	1.072142	1.022538	3.297588	.622253	.667567	-.6518	.909533

References

Bockstael, N. E., I.E. Strand and W. M. Hanemann. "Time and the Recreation Demand Model." *American Journal of Agricultural Economics* 69 (1987): 293-302.

Buse, A. "Evaluating the Linearized Almost Ideal Demand System." *American Journal of Agricultural Economics* 76 (1994): 781-793.

Cesario, F. J. and J.K. Knetsch. "Time Bias in Recreation Benefits Estimates." *Water Resources Research* 6 (1970): 700-704.

Chalfant, J. A. "A Globally Flexible, Almost Ideal Demand System." *Journal of Business and Economics Statistics* 5 (1987): 233-242.

Chalfant, J.A. and A.R. Gallant. "Estimating Substitution Elasticities with the Fourier Cost Function: Some Monte Carlo Results." *Journal of Econometrics* 28 (1985): 205-220.

Chen, H.Z. and A. Randall. "Semi-nonparametric Estimation of Binary Response Models with an Application to Natural Resource Valuation." *Journal of Econometrics* 76 (1997): 323-340.

Clawson, M. "Methods of Measuring the Demand for and Value of Outdoor Recreation." Washington, DC: Resources for the Future. 1959.

Creel, M. "Welfare Estimation Using the Fourier Form: Simulation Evidence for the Recreation Demand Case." *Review of Economics and Statistics* 79 (1997): 88-94.

Creel, M. and J.B. Loomis. "Semi-nonparametric Distribution Free Dichotomous Choice Contingent Valuation." *Journal of Environmental Economics and Management* 32 (1997): 341-358.

Deaton, A. and J. J. Muellbauer. "An Almost Ideal Demand System." *American Economic Review* 70 (1980): 312-326.

Diewert, W.E. "Intertemporal Consumer Theory and the Demand for Durables." *Econometrica* 42 (1974): 497-516.

Elbadawi, I., A.R. Gallant, and G. Souza. "An Elasticity Can be Estimated Consistently without a priori Knowledge of Functional Form." *Econometrica* 51 (1983): 1731-1751.

Gallant, A.R. "On the Bias in Flexible Functional Forms and an Essentially Unbiased Form: The Fourier Flexible Form." *Journal of Econometrics* 15 (1981): 211-45.

——— . "Unbiased Determination of Production Technologies." *Journal of Econometrics* 20 (1982): 283-323.

——— . "The Fourier Flexible Form." *American Journal of Agricultural Economics* 66 (1984): 204-208.

Larson, D.M. "Joint Recreation Choices and Implied Values of Time." *Land Economics* 69 (1993): 270-286.

Larson, D.M. and S.L. Shaikh. "Empirical Specification Considerations for Two-Constraint Models of Recreation Demand." Contributed Paper. American Agricultural Economics Association Annual Meeting. Toronto, Canada. 1997.

Larson, D.M., S.L. Shaikh and J.B. Loomis. "A Two-constraint AIDS Model of Recreation Demand and the Value of Leisure Time." Contributed Paper. XXII International Conference of Agricultural Economists. Sacramento, CA 1997.

McConnell, K.E. "Some Problems in Estimating the Demand for Outdoor Recreation." *American Journal of Agricultural Economics* 57 (1975): 330-39.

McConnell, K.E. and I.E. Strand. "Measuring the Cost of Time in Recreation Demand Analysis: An Application to Sportfishing." *American Journal of Agricultural Economics* 63 (1981): 153-156.

Piggott, N.E. "The Benefits and Costs of Generic Advertising of Agricultural Commodities." Ph.D. Dissertation. Dept. of Agricultural and Resource Economics, University of California at Davis. 1997.

Smith, T.P. "A Comparative Static Analysis of the Two-constraint Case." Appendix 4.1 in Benefit Analysis Using Indirect or Imputed Market Methods, Vol.2, N.E. Bockstael, W.M. Hanemann and I.E. Strand, eds. Washington, D.C.: Report to the Environmental Protection Agency, 1986.

Smith, V.K., W.H. Desvousges and M.P. McGivney. "The Opportunity Cost of Travel Time in Recreation Demand Models." *Land Economics* 59 (1983): 259-277.

White, H. "Using Least Squares to Approximate Unknown Regression Functions." *International Economic Review* 21 (1980): 149-70.

Visitor Day Equivalency Analysis: An Augmented Service-To-Service Methodology For Evaluating Recreation Site Quality Changes

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Abstract

We propose a methodology, Visitor Day Equivalency Analysis (VDEA), to estimate the magnitude of the compensation projects needed to compensate for the harm caused by lost recreation opportunities at a site. VDEA is consistent with NOAA guidelines for scaling compensatory restoration projects to off-set injury due to a release of a hazardous substance, since it uses a consistent approach for valuing service losses and gains. In the context of these guidelines, there are two basic types of scaling approaches: "value-to-value" and "service-to-service." In the former, one conducts economic valuation studies of the magnitude of damages *measured in dollars* and then looks for compensation activities which provide the same or greater *dollar* benefits. In the context of recreation, service-to-service scaling would be defined solely in terms of recreation visits. However, one might like to augment service-to-service studies with "weights" which allow more flexibility in scaling service flows. VDEA is an "augmented service-to-service" approach in the sense that it incorporates weights on visits to reflect relative marginal values over time and across damages and compensation projects.

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Introduction

In most situations where economists are called upon to value environmental services, budget is an issue. Thus, economists must economize on information acquisition. This is so especially when the event or policy being studied involves a relatively small effect and large-scale studies are deemed not worth the effort. Moreover, even if the subject action has a substantial economic impact, one is often called upon to provide an initial view of the potential magnitude of the effect as a first step in a more elaborate analysis.

Naturally, different valuation circumstances call for different strategies for conserving on information. The present paper is directed toward developing a benefits-transfer and/or value estimation method tailored to the following set of circumstances.

1. A release of some hazardous substance has caused injuries to natural resources which provide, *inter-alia,* recreation service flows.
2. The legal setting is such that resource-enhancement projects will be undertaken in order to provide in-kind compensation for the value of lost recreation use due to the release.
3. Data are relatively available on the amount of site visitation, including the number of recreational user days that occurred at the site pre and post the release.[1]
4. The number of user days that would be generated by the compensation projects can be estimated with a sufficient degree of precision to satisfy relevant stakeholders.
5. Data related to estimating the *changes in the value* of a site visit due to the injury and to the potential compensatory restoration projects are relatively reliable, but data related to the *total value* are relatively scarce and/or controversial.

[1] For convenience, we equate visitor days to trips. Adjustments for trips lasting two or more days may be necessary in some applications.

For example, suppose that a unique recreation area is closed by a spill. Managers of the site have data on the number of visitors during normal conditions. Negotiations are underway on potential compensation projects, and some scaling of them is needed. A reasonable estimate could be made of the effect of these projects on increases in the number of days to the site relative to baseline conditions. However, the value of a recreation day at the site is unknown, because no studies of this type of recreation have been undertaken previously. Moreover, because of the unique nature of the site, establishing the value of the site would be a complex, expensive undertaking, involving perhaps even a global set of origins and alternative destinations. What can be done?

We propose a methodology, Visitor Day Equivalency Analysis (VDEA), that may prove useful in such a circumstance. VDEA is used to estimate the magnitude of the compensation projects needed to compensate for the harm. VDEA is consistent with NOAA guidelines for scaling compensatory restoration projects, since it uses a consistent approach for valuing service losses and gains. In the context of these guidelines, there are two basic types of scaling approaches: "value-to-value" and "service-to-service." In the former, one conducts economic valuation studies of the magnitude of damages *measured in* dollars and then looks for compensation activities which provide the same or greater *dollar* benefits. The primary service-to-service scaling method, Habitat Equivalency Analysis (HEA) arises formally as an approximation to value-to-value measures (Jones and Pease). Given small projects with constant values across projects as well as over time, the marginal value measures for a change in environmental services drop out of both sides of the scaling equation, and one is left with changes in services. Of course, situations where all values are exactly (or even approximately) constant are rare, as least from an economist's perspective. Thus, one might like to augment service-to-service studies with "weights" which allow more flexibility in scaling service flows. The weights are, of course, necessarily economic.

In the context of recreation, service-to-service scaling would be defined solely in terms of recreation visits. VDEA is an "augmented service-to-service" approach in the sense that it incorporates weights on visits to reflect relative marginal values over time and across damages and compensation projects.

Theory

Suppose that there is a single representative individual, with a lifetime utility function that is additively separable across time periods which takes a quasi-linear form in each period. The indirect utility function for each period is defined by

$$(1) \qquad V(p,Q,Y) \equiv \text{Max } \{\lambda M + u(T,Q); \text{ s.t. } M + pT \leq Y\},$$

where λ is the marginal utility of the Hicksian composite M, which has a price of one, p is price of a trip to the site days, T is number of trips taken to the site, Q is a vector of characteristics of the experience, and Y is income. We can think of this function as arising from a two-staged budgeting problem. The present value of these V(.) functions might next be maximized by choice of income paths over time subject to a lifetime budget constraint. Naturally, if visitor data are by the number of days and not by trips, complications of trips of different durations must be addressed. Here, we assume for simplicity that all trips are one-day trips.

One might be curious as to why a quasi-linear form is specified. The reason is twofold. First, there is some concern in the valuation community about measuring willingness to accept compensation, as opposed to measuring willingness to pay. We avoid a choice if there are no income effects, and we can specify the value of both damages and compensation as being measured by willingness to pay. Second, and more importantly for our purposes, not having income effects in the trip-demand functions eliminates having to predict trips at different income levels (i.e., with and without Hicksian compensation being paid.) Existing data typically are the number of trips at full income, not the number of trips after payments to avoid environmental harm. Of course, it is the latter that matters, and, when income effects are present, one would need to know the demand function for trips. But if one has this demand function, then one can compute values directly, and our whole issue becomes moot. We are fairly comfortable with the assumption that there are no income effects, since recreation demand studies usually don't identify income as a key feature of trip demands (although exceptions to this statement do exist.)

The period-by-period welfare measures for damages are defined by,

(2) $$V(p,Q^0_t,Y - WTP_t) = V(p, Q^d_t,Y),$$

where Q^0_t is the baseline characteristics *but for* the injury and Q^d_t is the vector of characteristics with the release, and now we have Y as optimal wealth allocation to this date. Q^d_t includes any natural and augmented recovery. Note that baseline and damaged quality characteristics vary across time. If Q^d_t involves a site closure, $Q^d_t = 0$.

Let $W_t \equiv W(Q^0_t, 0, p, Y)$ be WTP_t if the site is closed. See Figure 1. If the site is not closed, let $w_t \equiv w(Q^0_t, Q^d_t, p, Y)$ be WTP_t. See Figure 2. Let δ_t be an indicator function, which equals 1, if there is a site closure in period t, and 0 if the site is open but damaged. Then, letting the discount factor be $\rho = 1/(1+r)$, the present value of damages is

(3) $$D \equiv \Sigma^B_t [\delta_t W_t + (1-\delta_t) w_t] \rho^t.$$

Here, we let *t* refer to time (in years), with *t*=0, being the time when the injury occurs and *t*=B, being the time when recreation services return to baseline.

Suppose that some compensation projects are to be undertaken. A project is a vector, A_t, of compensation activities. When these projects are implemented we obtain a vector of quality characteristics, $Q^c_t = Q(A_t)$. Per-period WTP for these projects is $\theta_t \equiv \theta(Q^d_{t,} Q^c(A_t))$ defined by

(4) $$V(p, Q^d_t,Y) = V(p,Q^C_t,Y - \theta_t).$$

The total benefits generated by these projects is,

(5) $\Sigma^{L}_{t=I}\theta_t\, \rho^t$

where, t=I is the time period when the increment in services from enhancement projects begin and, t=L is the time period when the increment in services from these projects end. See Figure 3.

Let C(A) be the cost of the compensating activities. A vector of activities, A, will compensate for the loss if (1) it is technically feasible and (2) satisfies the inequality

(6) $\Sigma^{L}_{t=I}\theta_t(Q^d_t, Q^c(A_t))\rho^t \geq D.$

Let the set of compensatory projects be Γ, i.e. Γ is the set of feasible projects A which satisfy (6). The objective of the regulator or responsible party is to find the least-cost compensatory projects, i.e. to identify

$$A^* \equiv \text{argmin} \{ C(A) \text{ s.t. } A \in \Gamma \}.$$

For convenience of presentation, we assume that the set of feasible activities is sufficiently rich that any cost-minimizing project will satisfy

(7) $\Sigma^{B}_{t} [\delta_t W_t + (1-\delta_t) w_t] \rho^t = \Sigma^{L}_{t=I}\theta_t\, \rho^t.$

If there are not many feasible projects, then it is unlikely that any one of them will *exactly* equate compensatory benefits with damages, but our argument goes through in more general settings using the obvious inequalities in place of (7).

We can rewrite this expression in terms of trips. Let T*(.) be the demand function for trips. We begin by defining $\Lambda_t \equiv W_t/T^*(p,Q^0_t,Y\text{-}W_t)$; Λ_t is the value per trip at the baseline characteristics. It is here that we make use of the assumption of quasi-linear preferences since $T^*(p,Q^0_t,Y\text{-}W_t) = T^*(p,Q^0_t,Y)$ in this case. Substituting for W_t in equation (7) gives,

(8) $\Sigma^{B}_{t} [\delta_t \Lambda_t T^*(p,Q^0_t) + (1-\delta_t) w_t] \rho^t = \Sigma^{L}_{t=I}\theta_t\, \rho^t.$

Dividing through both sides of (8) by Λ_0 (i.e, $W_0/T^*(p,Q^o_0)$) we get

(9) $\Sigma^{B}_{t}\,[\delta_t(\Lambda_t/\Lambda_0)T^{*}(\,p,Q^{0}_{t}) + (1\text{-}\delta_t)\,(w_t/W_0)T^{*}(\,p,Q^{0}_{0})]\;\rho^{t} =$

$\Sigma^{L}_{t=I}\,[(\theta_t\,/W_0)T^{*}(\,p,Q^{0}_{0})]\rho^{t}.$

Note that if there is a closure in the initial time period, the first term on the right hand side, T^{*}_{0}, is just the quantity of trips that would have occurred *but for* the incident. This is because $(\Lambda_t/\Lambda_0) = (\Lambda_0/\Lambda_0) = 1$. In subsequent time periods, if the closure continues, the damage can still be expressed in units of trips from the initial time period. However, these trips are weighted by the percentage change in welfare over time *but for* the incident (i.e., $(W_t/W_0)T^{*}_{0}$). Under certain conditions, this welfare change can be approximated by the change in demand over time.

Thus from a measurement perspective, one would need an estimate of the initial period's baseline visitor days *but for* the incident as well as how welfare (i.e., demand) is changing through time. The second term on the left hand side has a similar interpretation. If the site is open, but injured, the damage to the trip experience can be expressed in units of initial period trips, *but for* the injury. These trips are weighted by the percentage loss in welfare due to the change in quality attributes of the site, or in other words, the value of the quality changes relative to the full value of a trip. It is worth noting that we can replace (w_t/W_0) by $[(w_t/W_t) * (W_t/W_0)]$, where (w_t/W_t) is the share in value that is lost due to the injury and (W_t/W_0) is the change in base line value *but for* the incident. However, for measurement purposes, it is "only" necessary to estimate the loss in the value of a trip experience as a percentage of its full value *but for* the incident times the baseline number of trips.

Finally, the right hand side of equation (9) says that welfare generated by the compensation activities can be also be measured as a weighted sum of the trips from the initial period. Here the weights, (θ_t/W_0), are given by the enhancement to the value of a trip as a percentage of the full value of the unimpaired trip in the initial period. These weights can be replaced by $(\theta_t/W_0) = [(\theta_t/W_t) * (W_t/W_0)]$, which is the product of the share in value from compensation projects and the percentage change in value over time.

What we see from equation (7) is that both damages due to the injury and the value of compensation projects are measured in terms of the baseline numbers of trips to the site.

Discussion

VDEA serves to allow one to refocus debate away from the value of a trip to the site and toward relative changes in value. For example, one weight of importance in VDEA is the change in value of a trip over time. Recall that for the case of linear demand curves, consumer surplus is given by trips squared divided by the marginal value of a trip. Then, with a constant slope over time, the change in value in one period relative to the first period is just the ratio of the number of trips in that period squared to the squared number of trips in the first period. One might not have a good demand model for trips to a unique site, but demand projections based on historical data and other considerations might be available. This allows an important by-pass of the strict service-to-service assumption of constant value, but does not require going all the way to estimating demand curves and surplus.

A similar point can be made regarding the weights on relative values of injury-causing events and compensatory restoration activities. Imagine that one wants to find out about how a project at a recreation site, such as an interpretive program, augments the value of a visit. A wider array of research possibilities is admitted than estimating a demand curve. Benefits transfer studies might be available about changes in surplus that can be agreed upon more easily than a transfer of the surplus itself. And, an on-site study of recreators using, say, a conjoint-type format (or even qualitative research) might be employed to arrive at insight into (i.e. agreement upon in a negotiation) relative changes in surplus.

Certainly not all, or perhaps even many, valuation situations would point towards using VDEA. We would be the first to recognize it as no panacea. However, we have encountered circumstances where such an approach has proven to be useful. In our experience, the most compelling situations for using Visitor Day Equivalency Analysis involve scaling compensatory restoration projects aimed at enhancing the injured recreation site.

This augmented service-to-service methodology for evaluating recreation site quality changes expands the opportunities for applying similar methods for evaluating service losses and gains. Absent such a methodology, one may be compelled to abandon in-kind compensation in favor of pursuing monetary compensation for the value of lost recreation services. Ultimately, the monetary compensation would be used for in-kind compensatory restoration projects anyway. However, the end result may be to over-compensate or under-compensate the public, depending upon how the public values the resultant site quality changes. By conducting the analyses necessary for ensuring the equivalency of the value of service losses and gains during the process of scaling compensatory restoration, such under or over compensation can be avoided.

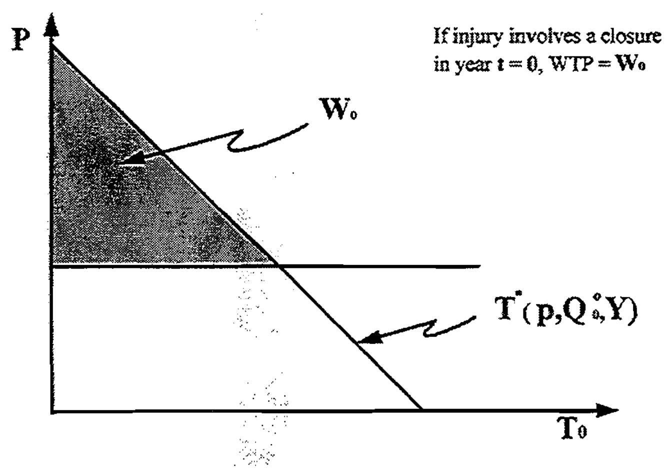

Figure 1. Baseline demand for Trips, T, in year t = 0.

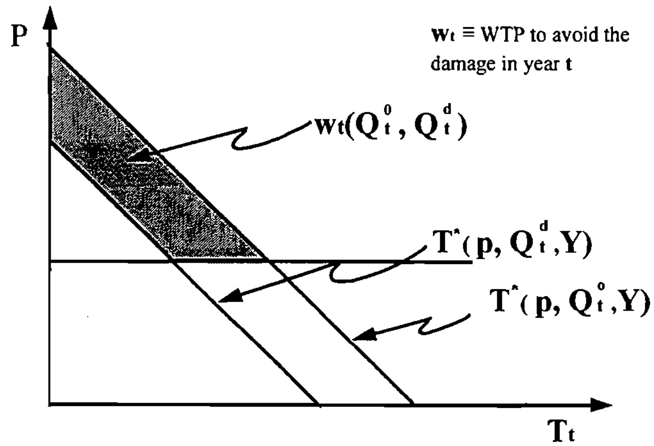

Figure 2. Demand shift due to injury.

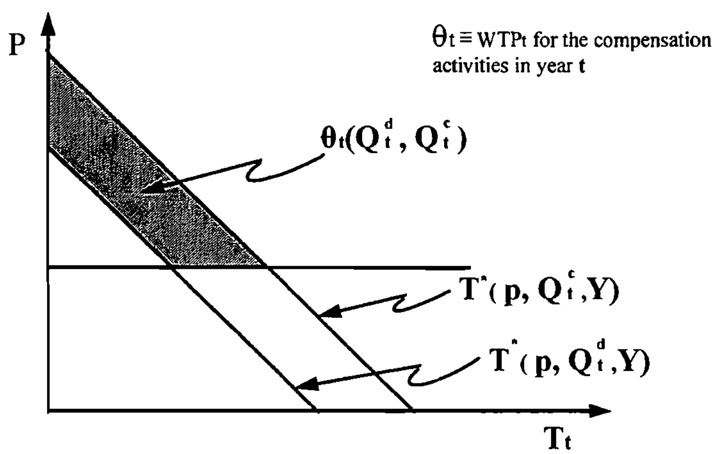

Figure 3. Demand shift due to compensation activities.

References

Jones, C.A. and K.A. Pease. *forthcoming*. "Restoration-Based Measures of Compensation in Natural Resource Liability Statutes," Forthcoming in *Contemporary Economic Policy.*

National Oceanic and Atmospheric Administration (NOAA). 1997. *Natural Resource Damage Assessment Guidance Document: Scaling Compensatory Restoration Actions,* Prepared by the Damage Assessment and Restoration Program, NOAA, 1305 East-West Highway, SSMC #4, Silver Spring, MD. 20910, December, 1997.

The Effect of Trip Lengths on Travel Cost Parameters in Recreation Demand

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Abstract

The random utility model has become a common framework for modeling and estimating recreational demand systems in the presence of a large number of substitute recreation sites. This paper uses the random utility model to examine the implications of distinguishing trips by trip length. Using data on fishing trips in Michigan, we find that the estimated travel cost parameter is significantly smaller for multiple day trips than for single day trips. The estimates suggest that multiple day trips account for a large share of the value of a change in the quality of recreation sites even when they are a small share of total trips. This paper addresses the modeling and welfare measurement issues that arise when such a finding is made.

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The Effect of Trip Lengths on Travel Cost Parameters in Recreation Demand

The random utility model is widely used to model the demand for recreation sites. The random utility model (RUM) offers a tractable way to include the prices and qualities of a large number of substitutes sites, as well as a way of linking demand to site quality variables. In addition, since the RUM is explicitly developed from budget constrained utility maximization, it permits one to readily derive utility theoretic welfare measures. Here we examine a RUM of recreational site choice that differentiates trips by trip lengths. Specifically, the model treats single and multiple day trips as distinct alternatives in the RUM choice set. In doing so, we find that the estimated travel cost parameters, if allowed to vary by trip lengths, are significantly different.

The empirical finding is important for environmental valuation since few recreational demand studies using RUM include multiple day trips, and even fewer model trip length choices (notable exceptions include Jones and Sung; Feather; Kaoru; and Shaw and Ozog). The statistical result suggests that multiple day trips are less responsive to changes in travel cost, and therefore, multiple day trips are relatively more valuable than single day trips. Thus, the finding suggests that a large share of the value for a change in environmental quality may be attributable to multiple day trips even when these trips constitute a small share of all recreation trips.

While RUMs offer a utility theoretic link that can be utilized to define and measure welfare changes, not all statistical models are consistent with this linkage. In the case we discuss, having separate price parameters for different trip lengths is not consistent with the usual derivation of a budget constrained RUM. Such a finding places the analyst in an awkward position. How should the statistical model be interpreted? Should a simpler model be imposed on the data, or should one search for a more complex utility model that might be more consistent with the statistical results? Rather than estimate a variety of more complex models, we take a different course. We examine alternative welfare measurement concepts such as consumer surplus, to see what they yield in the model with separate travel cost parameters for the different trip lengths. These approaches are not presented as definitive answers. Rather, we consider this work to be a first step in the investigation of these issues. The overriding result is that the fairly common practice of ignoring multiple day trips may substantially bias values for changes in environmental quality.

The paper is organized as follows: first, we provide a brief overview of the RUM and explain how it can be used to derive welfare measures. We then present and discuss consumer surplus and some other welfare measurement concepts that could be applied to our empirical model that has different travel cost parameters for single and multiple day trips. Because the empirical model bears some resemblance to a model that is non-linear in income, we also briefly RUMs with income effects. The next section of the paper presents our empirical example. Again, the main finding was that there were significant differences in the travel cost parameters for trips of different lengths. We then apply some alternative measurement concepts to the estimated models and discuss the results.

The Random Utility Model (RUM)

In order to motivate some of the issues related to the implications and complications of our empirical findings, we briefly review of the theory of discrete-choice known as the random utility model with a focus on recreational site choice.[1] The following notation will be used throughout:

Y	Income
k,j	represent sites
z	Hicksian composite good
p_k	price of a trip to site k (travel cost)
δ_k	indicator function: equals one if site k is chosen, equals 0 otherwise.
q_k	quality characteristics at site k
$U(\cdot)$	utility function
$V(\cdot)$	deterministic part of utility (from researcher's perspective)
ε_k	stochastic part of utility for site k (from researcher's perspective)

[1] In doing so, we borow heavily from the literature on RUMs (McFadden, 1981; Small and Rosen; Hanemann 1982; Ben-Akiva and Lerman; Train). Presentations of the RUM that are geared toward recreational demand are given in Bockstael, Hanemann, and Strand; Bockstael, McConnell and Strand; and Freeman. Econometric methods for estimating RUM models have been reviewed by McFadden 1981; and more recently by Morey 1998.

λ_k Marginal utility of income conditional on choice of k

Using the above notation, the decision of which site to visit on any choice occasion can be written as follows:

$$\max_{\{z,\,\delta\}} \; U(z, q) \tag{1}$$
$$\text{s.t.} \quad Y = \Sigma_k \, p_k \, \delta_k + z$$
$$\Sigma_k \, \delta_k = 1, \;\; \Sigma_k \, \delta_k q_k = q, \;\; \delta_k \in \{0,1\}.$$

The first constraint is the budget constraint. The indicator function, δ, reflects the condition that only one alternative can be chosen, and this notion is embodied in the second row of constraints. The quality constraint indicates that the quality variable entering the utility function is the quality of the site actually chosen -- site quality and site choice are weak complements. The model formulation states that the individual receives utility from a composite good and from the characteristics of a single discrete alternative.

Budget exhaustion implies that $z = Y - \Sigma_k \, p_k \delta_k$, i.e., any income not spent on the chosen alternative is spent on the composite good. Hence, conditional on selecting site k, the budget constraint can be restated as $z = Y - p_k$. Substituting for z in U(·), the conditional indirect utility for the choice occasion (maximum utility conditional on choosing k), is $U_k = U_k\,(Y\text{-}p_k\,,\, q_k\,)$. The problem in (1) can now be stated as choosing the site k that maximizes the conditional indirect utility U_k, that is,

$$\text{Max}_{\{k\}} \;\; U_k\,(Y\text{-}p_k\,,\, q_k\,). \tag{2}$$

Based on (2), when one observes an individual choosing site k' from a set of K alternatives, then by revealed preference it can be inferred that $U_{k'} = \max\{U_1\,, U_2\,,..., U_K\,\}$ for that individual. Specifically, the choice of k' implies the following set of utility inequalities: $U_{k'} > U_k$ for all $k \neq k'$. The model becomes the "random" utility model by recognizing that not all attributes that affect utility can be observed by analysts. Thus, analysts can only predict the probability that any alternative is the best in the choice set as follows:

$$\pi_k \;=\; \text{Prob}[\delta_k{=}1] \;\;=\;\; \Pr\,[U_k = \max\{U_1\,,..., U_K\,\}] \tag{3}$$

$$= \Pr [\, U_k - U_1 > 0, \ldots, U_k - U_{k-1} > 0, U_k - U_{k+1} > 0, \ldots, U_k - U_K > 0 \,].$$

The choice probabilities serve as the *expected* demand functions for each alternative in the choice set.

In the RUM, utility is typically given by the sum of a deterministic portion of utility that can be measured (estimated) by analysts and a stochastic portion that is unobservable for analysts but remains known to the individual making the choices. That is, conditional utility is

$$U_k = V_k + \varepsilon_k, \tag{4}$$

where V is the deterministic portion of utility to be estimated. For example, researchers commonly adopt a linear form such as

$$V_k = \lambda(Y - p_k) + \beta q_k. \tag{5}$$

Here the λ is the marginal utility of money as well as the negative of the coefficient on price, and β is a vector of marginal utilities associated with a vector of site characteristics q_k. Alternatively, a non-linear-in-income form can be specified,

$$V_k = f(Y - p_k; \lambda) + \beta q_k. \tag{6}$$

but notice that income and price will still enter as a single term, $Y - p_k$, because of budget clearing and substitution though the composite good.[2] A particularly simple form of f(·) would be

$$V_k = \lambda_k(Y - p_k) + \beta q_k. \tag{7}$$

In this case, the income variable will not cancel from the utility differences as it does in (3) when $\lambda_k = \lambda$ for all k. In (7), the marginal utility of money (or equivalently, the price responsiveness given by the travel cost parameter) is dependent on the alternative chosen, i.e., it is state dependent.

To derive an econometric model, one assumes a distribution for the error terms. If the errors are assumed to be drawn from a generalized extreme value (GEV) distribution (McFadden, 1978) then the resulting site choice probabilities (3) have closed forms. A simple form of GEV distribution is the Type 1 Extreme Value (EV) distribution. To simplify what follows, discussion will focus on EV models.

[2] Note that RUMs with income effects require knowledge and definition of choice occasion income. Lacking a model of how income is allocated to choice occasions, choice occasion income becomes another variable that must be chosen by the researcher, and, due to the income effect, the estimated model will be sensitive to the definition of choice occasion income.

When the errors are assumed to be i.i.d. with an EV distribution, choice probabilities will have the familiar multinomial logit (MNL) form,

$$\pi_j = e^{V_j} / \Sigma_j e^{V_j}. \tag{8}$$

From the above expression, it is clear that the π_j will depend not only on the price and quality of site j, but on the prices and qualities of all sites. Therefore, the RUM model provides a convenient form for relating the demand for a site to the prices and qualities of all relevant substitute sites.

What models are estimated here? The models estimated here have indirect utility functions of that take the following form

$$V_{k,i} = -\lambda_i p_k + \beta q_k. \tag{9}$$

where i represents single day (sd) and multiple day (md) trips. In the first model we estimate, $\lambda_i = \lambda$ for both single and multiple day trips. A model with this constraint is consistent with the budget constrained RUM with a linear-in-income form that is constant across all alternatives, as reflected in (5). Alternatively, when the price parameter is allowed to differ across groups of alternatives, such as single and multiple day trips, we see no obvious way to cast this in the budget-constrained RUM framework. This begs the question of what to do with such a model in terms of welfare measurement, and we address this issue in the next section.

Now suppose one estimates a MNL and lets the price parameters differ across sub-groups of alternatives as in (9). One interpretation of such a model is that it resembles the MNL based on (7). We note here that there is an important difference between the two. If (7) represents the true model and instead one estimates (9), there is an omitted variable. To see this, redefine λ_k so that $\lambda_k = \lambda + \Delta_k$. Substituting this into (7), the deterministic part of the indirect utility is

$$V_k = \lambda Y - \lambda p_k + \Delta_k(Y - p_k) + \beta q_k. \tag{10}$$

Here, the λY will cancel across all alternatives, and the $\Delta_k(Y-p_k)$ will be an omitted variable if the form in (5) is imposed. During model estimation, if the income is omitted and the alternatives are given their own price parameters, as in (9), then the only omitted variable is $\Delta_k Y$.[3] We are not suggesting that (10) is the "correct" model. Rather we discuss it because it resembles the model based on (9) which is what we estimated in the empirical section.

Welfare Measurement and Consumer Surplus

It is well known that the welfare measures for RUMs based on EV errors can be defined using the inclusive value concept. In particular, with the EV errors and linear in income indirect utility functions, the compensating variation for a general multiple site change in quality (from 0 to 1) is given by

$$CV = \{\ln[\Sigma_j \exp(V_j^1)] - \ln[\Sigma_j \exp(V_j^0)]\}/\lambda = \{IV^1 - IV^0\}/\lambda \qquad (11)$$

where IV represents the inclusive value. This measure is suitable for evaluating the addition or elimination of sites, as well as measuring the effects of changes in site quality at multiple sites. Equation (11) is the appropriate welfare measure for models with indirect utility given by (5).

Another useful result for RUMs with EV and GEV errors is that $\partial IV/\partial V_k = \pi_k$ (McFadden, 1981). Using this result, one can see that when the model is linear in income, calculating the area behind the demand curve for any site, the integral of π_k from the current price to infinity, returns the welfare measure for access to that site. When the model is linear in income, this consumer surplus measure is exactly the same as the equivalent or compensation variation. One can also use this to establish that for changes in quality, weak complementarity holds in the RUM.

Non-linear in income models (some bounds): In the case of non-linear income effects, McFadden (1997) has shown that the welfare measurement requires a more complex approach that does not yield a closed form solution (see Herriges and Kling for an application of the approach). McFadden

[3] This formulation suggests possibilities for how the omitted variable might affect estimation if (7) were the true form. For example, if $\Delta_k Y$ is distributed EV over individuals, it can be incorporated into the error term and will manifest as a larger variance on some alternatives. This effect might be captured by a nesting structure. We note that model versions that nested the alternatives by trip lengths did not resolve the differences in the travel cost parmeters by trip lengths (see Hoehn et al).

also presents some bounds on the welfare measure for a non-linear model that we will also apply here. These are essentially an application of the marginal welfare measures presented by Hanemann (1983). In the case where the indirect utility functions take the form in (7), the bounds are given by

$$\Sigma_k \ \pi_k{}^0 \, \Delta q_k \, \beta/\lambda_k \ \leq \ \text{true measure} \ \leq \ \Sigma_k \ \pi_k{}^1 \, \Delta q_k \, \beta/\lambda_k \qquad (12)$$

We will make use of these bounds later. Another method for approximating the welfare measure in the non-linear case has been given by Morey (1998) -- see also Morey, Rowe, and Watson. When the RUM is specified with income effects, two properties that affect all demand models with income effects are worth note. First, the consumer surplus measure will no longer equal the Hicksian measures. Second, the consumer surplus measure for changes in the prices or qualities of multiple goods (sites) will no longer be path independent.

Consumer surplus when λs vary by alternatives: Now, further consider the case where $\lambda_k \neq \lambda_j$ as in the indirect utility functions in (7) and (9). In the case of (7), the model is formally a non-linear in income model and compensating and equivalent variation measures could be defined following McFadden (as mentioned above). However, for (9), there is no explicit link between the choice probabilities and income which make definition of Hicksian surplus measures difficult. None-the-less, we can still define consumer surplus for this model (as well, weak complementarity measures based on consumer surplus can be defined). What do consumer surplus measures look like in this case? If there is only a change in quality at a single site, k, then the change in consumer surplus is given by the change in area behind π_k.[4] With GEV or EV errors, the change in consumer (CS) for a change in site k quality ($q_k{}^0 \rightarrow q_k{}^1$) is

$$CS_k{}^1 - CS_k{}^0 = \{IV(q_k{}^1) - IV(q_k{}^0)\}/\lambda_k \qquad (13)$$

where $CS_k{}^0$, $CS_k{}^1$ represent the consumer surplus for alternative k under the quality levels $q_k{}^0$ and $q_k{}^1$. Because of the non-linearity of income, for changes in quality at *multiple sites*, weak complimentarity

[4] We try to refer to this as the "consumer surplus based weak complimentarity measure" to distinguish it from Hicksian surplus measures, though we will sometimes just refer to it as weak complementarity.

measures based on consumer surplus will depend on the path or sequence of quality changes. (For a general discussion of weak complementarity with multiple site changes in quality, see Bockstael and Kling). When $\lambda_k=\lambda$ for all k, then for multiple site changes in quality, the weak complimentarity measure based on consumer surplus changes is path independent -- it reduces to a comparison of IV where all sites are evaluated at their initial levels of quality and IV where all sites are evaluated at their post-policy level as in (11).

In the case we are interested in, the travel cost parameter is allowed to vary by single and multiple day trips, but remains constant within these two trip lengths. Hence, the weak complementarity measure for changes in quality at multiple sites within a trip length is independent of the sequence in which site quality is changed. However, when one evaluates changes across trip lengths, the measure is no longer path independent because the denominators (λ_k) place differing weight on the utility changes. Using the result in equation (12), one can collapse all quality changes for single day alternatives into one expression (likewise for multiple day trips). Thus, we'll let the expression, $CS_{sd}(q_{sd}^1)$-$CS_{sd}(q_{sd}^0)$, represent the weak complementarity measure of changes in the consumer surplus for single day trips where q_{sd}^0 and q_{sd}^1 represents a vector of changes in quality at single day alternatives. Note that if single and multiple day trips are substitutes, then these CS measures will depend on the level of quality for multiple day sites. Moreover, when the alternatives are indexed by trip lengths as well as site locations, it is not physically possible to change quality at some location without changing quality at multiple alternatives in individuals' choice sets. Thus, any change in quality will affect multiple alternatives in the choice set making it necessary to worry about the path.

With this in mind, it is convenient to define the consumer surplus based weak complementarity measure according to two possible paths of quality changes. Along the first path, the changes in quality are evaluated at the single day sites first, and then the changes in quality are evaluated at the multiple day sites given the new level of quality at single day sites.

Define this as

$$WC(sd,md) = CS_{sd}(q_{sd}^1, q_{md}^0) - CS_{sd}(q_{sd}^0, q_{md}^0) + CS_{md}(q_{sd}^1, q_{md}^1) - CS_{md}(q_{sd}^1, q_{md}^0)$$

$$= \{IV^{10} - IV^{00}\}/\lambda_{sd} + \{IV^{11} - IV^{10}\}/\lambda_{md} \quad (14)$$

where WC(sd,md) stands for the weak complementarity defined by changing single day (sd) site quality first, and then evaluating the change in multiple day (md) site quality given the new level of single day site quality. In the IV expressions, the superscript ij represents the single and multiple day quality vectors. For example, IV^{10} represents the inclusive value evaluated at (q_{sd}^1, q_{md}^0).

Along the second path, the changes in quality are evaluated at the multiple day sites first, and then the changes in quality are evaluated at the single day sites given the new level of quality at multiple day sites:

$$WC(md,sd) = \{IV^{10} - IV^{00}\}/\lambda_{sd} + \{IV^{11} - IV^{10}\}/\lambda_{md} \quad (15)$$

If the travel cost parameters were the same across trip lengths, these measures collapse into the familiar changes in inclusive value. In a later section, these two paths will be applied to the empirical model that has differing travel cost parameters for each trip length.[5]

<u>An approach resembling separate models</u>: Now suppose one were to have estimated separate models. Estimating separate models for single and multiple day trips is akin to allowing the parameters to differ across models and eliminating all substitution across trip lengths. This then suggests another possible way to measure changes in welfare from the jointly estimated model. This approach would constrain the predicted total single day and multiple day trips so that they are held at their initial levels, π_{sd}^0 and π_{md}^0 (which would have been the case if separate models were run). Welfare measures can then be defined by conditioning them on a particular trip length. For changes in quality, define

[5] These are two among many possible paths. For example, along another path one could calculate the change in surplus for a change in $q_{1,sd}$; then do it for $q_{1,md}$; then for $q_{2,sd}$; $q_{2,md}$; $q_{3,sd}$; $q_{3,md}$; etc. However, the above two paths are the most extreme. The intuition for this is that because of substitution effects, smaller (larger) shifts in multiple day trip demand occur when single day trips are evaluated at higher (lower) levels of quality.

$WC^0_{conditional}$ as the weak complementarity measure based on the conditional on a trip length welfare measures, as follows:

$$WC^0_{conditional} = \pi_{sd}{}^0 CV_{sd} + \pi_{md}{}^0 CV_{md}$$

$$\text{and} \tag{16}$$

$$WC^1_{conditional} = \pi_{sd}{}^1 CV_{sd} + \pi_{md}{}^1 CV_{md}$$

where $CV_i = (IV_i^1 - IV_i^0)/\lambda_i$ for i = sd, md. The conditional on a trip length inclusive values are defined by $IV_i = \ln\{\Sigma_{j\epsilon i}\, e^{V_j}\}$ where the summation is over trips of length i. In the model with the indirect utility function in (7) with linear but differing travel cost parameters, these measures would bound the "true" welfare measure just as McFadden's bounds would. The only difference is that this measure uses more information because it makes use of the fact that the income effects are constant over some alternatives. Note that these are bounds on the welfare measure that would be derived had the model using (7) as the indirect utility been estimated. The difference in our case is that model (9) is estimated, not (7). In this case the measure in (16) is akin to what one would get from separate models for the different trip lengths.

Empirical Example

Two models of recreational fishing at the Great Lakes are estimated to illustrate how the travel cost parameter can differ across trip lengths. The models are specified as multinomial logits with two trip lengths: single and multiple day trips. The first model is specified with a single price parameter across both trip lengths, and the second model allows the trip length parameter to vary across the single and multiple day trip lengths. In order to focus on the effect of trip lengths on the travel cost parameters, the complexity of the models is kept to a minimum.

The models use a sample of Michigan anglers who took Great Lakes trout and salmon fishing trips in 1994. The data are drawn from a larger survey of Michigan anglers (see Hoehn et al for a complete description of the survey). There are 312 trips of which 78% are single day trips and 22% are multiple day trips. The sites are defined as the stretch of shoreline within any of Michigan's counties that border Lakes Michigan, Huron, and Superior. The two models are specified as multinomial logits with indirect utility taking the forms in (5) and (9). For single day trips, only sites within 150 miles of an anglers residence are feasible. For multiple day trips, all the sites are feasible. Hence, sites can appear in the choice set up to two times: once as a destination for a single day trip and once as a destination for a multiple day trip. Each site is characterized by its catch rate, travel cost, and dummy variables for the trip length and for the lake it borders. Catch rates are derived from negative binomial models of catch per hour which were developed in an analysis of a separate creel survey data set (Lupi et al, 1998). The catch rates are specific to sites and vary on a monthly basis. Travel costs are defined as the round trip driving costs plus lodging and time costs (Hoehn et al.).

The parameter estimates for the two models are reported in Table 1. The table also presents the log-likelihood values for each model. As a statistical model explaining the data, the version with the single price parameter is overwhelmingly rejected in favor of the model with separate price parameters for each trip length. We note that the result that the multiple day trips were much less responsive to price than single day trips was consistent across a number of model specifications.[6] The result is consistent with the findings of Jones and Sung and the results of our earlier work with the data (Hoehn et al.) even though both of these efforts involved models with much broader ranges of fishing types and sites as well as complex multi-level nesting structures. The implication is that multiple day trips are relatively more valuable than single day trips.

Interestingly, the estimated travel cost parameter for the model 1 is closer to zero than either travel cost parameter for model 2. We would have expected the constrained parameter to lie between the unconstrained parameters. Of course, due to the discrete nature of the data, the individual parameters are

[6] For example, alternative specification of travel costs based on different approaches for the value of travel time, did not resolve the issue. Changing the definition of travel costs did affect the parameter estimates. However, regardless of the definition of travel costs and time value, all models exhibited a significantly better fit when single and multiple day trips were allowed to have different travel cost parameters.

of less interest than the ratio of the parameters. In the model with a single price parameter (model 1), the ratio of the catch rate parameter to the negative of the travel cost parameter, β/λ, gives the value of an across the board one unit change in catch rates. In the table, comparable measures are defined within each trip type for model 2 (that is, β/λ_{sd} and β/λ_{md}). Table 1 also reports the weighted average of these where the weights used are the shares of trips (78% single day and 22% multi-day). The marginal valuation for model 1 is larger than the marginal value for either trip length in model 2, and it is over two and a half times the weighted average for model 2. Moreover, over half the value of an across the board marginal change in catch rates is due to multiple day trips even though multiple day trips represent less than one fourth of the trips in the sample.

Non-marginal welfare measures: The above sections have discussed some of the issues involved in attempting to define and derive welfare measures for models based on indirect utility functions taking the form in (9). In light of this, we apply several alternative approaches to the two multinomial logit models we have estimated. The approaches are as follows.

1. Ignore the statistical evidence and impose a single travel cost parameter (use model 1).
2. Pick one of the travel cost parameters and use it in equation (11) as if it were the unique estimate of the marginal utility of money. This leads to two approaches since we have two different travel cost parameters.
3. A variant of 2. would use the weighted average of the two travel cost parameters and treat the average parameter as the unique estimate of the marginal utility of money for use in (11). The weights are the initial shares of trips for each trip length. This approach has been used by some when sequential estimation was applied to nested logits, since with sequential estimation parameters across nests cannot easily be constrained (Jones and Sung).
4. Select a path, and calculate the consumer surplus measure. In our case, we will consider two possible paths. The first involves changing site quality at sites in the single day branch of the model and then changing site quality at sites in the multiple day branch of the model. The

second would change quality at multiple day sites and then change quality at single day sites.

5. Treat the model as an approximation to a non-linear-in-income model and apply McFadden's bounds that were developed for models that are non-linear in income.

6. Calculate the sum over different trip types of the trip weighted conditional welfare measures as given in (16) above.

In order to illustrate how each of these alternative measurement concepts compare to one another, each was calculated for a hypothetical doubling of the catch rates at Lake Huron. The results are reported in Table 2. The per-trip values in table 2 represent sample averages.

The results of the different measurement concepts reveal several insights. First, the consumer surplus value of about $24 per trip that is derived from model 1 is the largest per-trip estimate of any in Table 2. This was expected given the large marginal value of catch changes reflected in the parameter estimates for model 1. For model 2, there is a substantial spread between the consumer surplus estimates depending on the path that is used to calculate surplus. Since $\lambda_{md} < \lambda_{sd}$, changing the quality of multiple day sites while holding single day sites at initial quality levels leads to a larger surplus estimate than the converse. The McFadden bounds, when applied to model 2, resemble an extrapolation from the site and trip specific marginal values. Since the doubling of Lake Huron catch rates results in some large site probability changes to some of the Lake Huron sites, the bounds are not very tight. In Table 2 they differ by a factor of three (recall the application of these bounds to model 2 is somewhat ad hoc since they are strictly developed for a model with indirect utility as in (7)). Interestingly, the use of either λ_{sd} or λ_{avg} in the conventional inclusive value welfare measure results in welfare measures that lie below the McFadden lower bounds. Again, the McFadden lower bounds are the extrapolation the marginal values using the probabilities in the initial state. This suggests that the measures based on either λ_{sd} or λ_{avg} are unreasonably low. As expected, the conditional on a trip length measures are tighter than the McFadden bounds.

Unfortunately, the comparisons presented in Table 2 do not yield a clear picture of what one should do to measure values in a model such as model 2. In part this is due to the fairly significant dependence of the consumer surplus measures on the path of the quality changes. The sensitivity of the consumer surplus based measures to the path (as well as the spread between the conditional on a trip length welfare measures) is directly related the degree of substitution among the trip lengths. The multinomial logit models estimated here were selected to simplify the trip length issue. However, the simple multinomial logits impose a high degree of substitution across trip lengths. Conversely, in a more extensive model of fishing in Michigan that used a broader range of fishing alternatives and that had trips nested by trip lengths, Lupi et al (1997) find that there was very little estimated substitution across trip lengths. Consequently, in the model of Lupi et al. (1997) the difference between the conditional on a trip length measures in (16) was typically less than 1 percent. In contrast, applying the McFadden bounds (or an ad hoc choice of λ_{sd} or λ_{md}) to the Lupi et al. (1997) model would produce measures as divergent as those for the simple model presented here.

Before closing we note that the idea that alternative trip lengths might differ in their price responsiveness would not be unusual or problematic in most economic demand systems. However, this is a potential problem in the RUM because of the constraint on the way income and travel costs enter the indirect utility function. Thus, to some extent, the results presented here indicate the need for more flexible demand models that are consistent with the statistical resutls (differing price responsiveness for different trip lengths) but that can also do what the RUM can (namely, link demands to site quality and accommodate a large number of substitutes in a tractable manner). This need is further emphasized when the statistical results of model 2 are taken at face value: a marginal change in catch rates is over four times as valuable for multiple day trips than for single day trips. Put differently, over half the value of an across the board marginal change in catch rates was due to multiple day trips even though multiple day trips represent less than one fourth of the trips in the sample. This suggests that substantial use values may be left unaccounted for by the common practice of dropping multiple day trips from recreational demand analyses.

References

Ben-Akiva, Moshe, and Steven R. Lerman, *Discrete Choice Analysis: Theory and Application to Travel Demand*, Cambridge: MIT Press, 1985.

Bockstael, Nancy E., Kenneth E. McConnell, and Ivar E. Strand, Jr., "Recreation," in: *Measuring the Demand for Environmental Quality*, (Braden, John B., and Charles D. Kolstad, eds.), New York: North-Holland Publishing, 227-355, 1991.

Bockstael, Nancy E., W. M. Hanemann, and I. E. Strand, "Measuring the Benefits of Water Quality Improvements Using Recreation Demand Models," Vol. II of *Benefit Analysis Using Indirect or Imputed Market Methods*, EPA Contract No. CR-811043-01-0, 1984.

Bockstael, Nancy E., and C. L. Kling, "Valuing Environmental Quality: Weak Complimentarity with Sets of Goods," *American Journal of Agricultural Economics*, 654-662, August, 1988.

Feather, Peter Milton, *Valuing Water Quality Using Discrete Choice Models*, Ph.D Dissertation, Department of Agricultural and Applied Economics, University of Minnesota, January, 1992.

Freeman III, A. Myrick, *The Measurement of Environmental and Resource Values: Theory and Methods*, Washington, D.C.: Resources for the Future, 1993.

Hanemann, W. M., "Marginal Welfare Measures for Discrete Choice Models," *Economic Letters*, 13:129-36, 1983.

Hanemann, W. Michael, "Applied Welfare Analysis with Qualitative Response Models," Working Paper # 241, University of California - Berkeley, October, 1982 .

Herriges, Joseph A., and Catherine L. Kling, "Nonlinear Income Effects In Random Utility Models," Working Paper, Department of Economics, Iowa State University, 1996.

Hoehn, John P., Theodore Tomasi, Frank Lupi, and Heng Z. Chen, *An Economic Model for Valuing Recreational Angling Resources In Michigan*, Report Submitted to the Michigan Department of Natural Resources and the Michigan Department Environmental Quality, December, 1996.

Jones, Carol Adair, and Yusen D. Sung, *Valuation of Environmental Quality at Michigan Recreational Sites: Methodological Issues and Policy Applications*, Final Report, EPA Contract No. CR-816247-01-2, September, 1993.

Kaoru, Yoshiki, "Measuring Marine Recreation Benefits of Water Quality Improvement by the Nested Random Utility Model," *Resources and Energy Economics*, 17(August): 119-136, 1995.

Lupi, Frank, John P. Hoehn, Heng Zhang Chen and Theodore Tomasi, "The Michigan Recreational Angling Demand Model," *Agricultural Economics Staff Paper* 97-58, Michigan State University, December, 1997.

Lupi, Frank, John P. Hoehn, and Douglas B. Jester, *Trout and Salmon Catch Rates at Michigan Great Lakes Sites, 1986 to 1995*, report to the Great Lakes Fishery Commission, March, 1998.

McFadden, Daniel, "Computing Willingness-To-Pay in Random Utility Models," Working Paper, University of California, 1997.

McFadden, Daniel, "Econometric Models of Probabilistic Choice Models," in: *Structural Analysis of Discrete Data with Applications*, (Manski, Charles, and Daniel McFadden, eds.), Cambridge: MIT Press, 1981.

Morey, Edward R., *TWO RUMs un CLOAKED: Nested-Logit Models of Site Choice and Nested-Logit Models of Participation and Site Choice*, forthcoming in: *Valuing the Environment Using Recreation Demand Models* (C.L. Kling and J. Herriges, eds.) Edward Elgar Publishing, 1998.

Morey, Edward R., Robert D. Rowe, Michael Watson, "A Repeated Nested Logit Model of Atlantic Salmon Fishing," *American Journal of Agricultural Economics*, 75(3):578-592, August, 1993.

Shaw, W.D., and M. Ozog, Modeling Overnight Trips: An Apllication of the Repeated Nested Multinomial Logit Model, forthcoming *Environmental and Resource Economics*.

Small, Kenneth A., Harvey S. Rosen, "Applied Welfare Economics With Discrete Choice Models," *Econometrica*, 49(1):105-130, January, 1981.

Train, Kenneth, *Qualitative Choice Analysis*, London, England, 1986.

Table 1: Two RUMs, with travel cost parameters that do and do not vary by trip length.

Model 1 ($\lambda_{sd} = \lambda_{md}$)		Model 2 ($\lambda_{sd} \neq \lambda_{md}$)	
Variable	Parameter (t-stat)	Variable	Parameter (t-stat)
Travel Costs	-0.013 (-9.71)	$TCost_{single\ day}$	-0.129 (-18.2)
		$TCost_{multi\text{-}day}$	-0.031 (-15.9)
Catch Rate	3.63 (4.53)	Catch Rate	7.60 (6.72)
Multi-day constant	2.98 (5.51)	Multi-day constant	-5.77 (-7.58)
Lake Superior constant	0.34 (1.18)	Lake Superior constant	2.03 (3.73)
Lake Michigan constant	0.44 (2.53)	Lake Michigan constant	2.01 (7.18)
Log Likelihood Value	-850	Log Likelihood Value	-516
Marginal Implicit Price		Marginal Implicit Price	
$\beta_{catch\ rate}/\lambda$	\$270.90	$\beta_{catch\ rate}/\lambda_{single\ day}$	\$59.10
		$\beta_{catch\ rate}/\lambda_{multi\text{-}day}$	\$242.81
		Weighted average β/λ	\$99.52

Table 2: The various measures for a doubling of Lake Huron catch rates.

Model and Measurement Method	Results		
	$	$	$
Model with $\lambda_{sd} = \lambda_{md}$			
Consumer Surplus		24.25	
Model with $\lambda_{sd} \neq \lambda_{md}$			
	Path: single day then multi-day		Path: multi-day then single day
Consumer Surplus	7.06		12.76
	Lower		Upper
McFadden bds.	5.98		18.47
Pick a λ, use usual formula	using λ_{sd}	using $\lambda_{avg.}$	using λ_{md}
$\{IV^1 - IV^0\}/\lambda$	3.77	4.54	15.48
Conditional CS	Lower		Upper
(single & multi-day trips held constant)	10.30		11.62

The Effect of Nesting Structure Specification on Welfare Estimation in Random Utility Models: An Application to the Demand for Recreational Fishing

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1. Introduction

The use of discrete choice models to estimate the welfare effects on recreators of a change in environmental quality is well established. Multinomial logit models are used in the recreation demand literature to model an individual's choice among qualitatively different recreation sites (see, for example, Adamowicz, 1994; Bockstael, Hanemann, and Kling, 1987; Feather, 1994; Jones and Sung, 1991; Karou, Smith and Liu, 1995; McConnell and Strand, 1994; Morey, Rowe and Watson, 1993; and, Parsons and Kealy, 1992). The model has been shown by Hanemann (1978, 1982) to be utility theoretic and capable of generating measures of compensating variation for changes in site quality.

An important drawback of the non-nested multinomial logit model is its reliance on the assumption of the independence of irrelevant alternatives (IIA). The IIA property imposes the restriction that the relative odds of choosing between two alternatives remains unchanged when a third alternative is introduced into the choice set; that is, the disturbances in the random utilities associated with the alternatives are uncorrelated. As a result, a rich variance in substitution patterns among sites that is most likely present in the data is assumed away. Economic theory holds that substitution among alternatives plays a pivotal role in welfare analysis; therefore, ignoring such substitution effects may result in erroneous estimates of welfare change.

There are two general ways to overcome the restrictiveness of the IIA assumption imposed by the logit model. The first is the use of multinomial probit (MNP) that allows correlation among the errors of different alternative-specific utilities to vary across every pair of alternatives. Therefore, MNP captures a rich correlation among alternatives; however, estimation requires multiple integration and can become computationally burdensome when the number of alternatives is greater than three. Simulated probabilities techniques have been developed to overcome this problem. These techniques are still in the early stages of development and applications have not, as of yet, been widely applied to recreation demand studies. (See Chen and Coslett (1998) for an application to recreation demand and Stern (1997) for a recent review of simulation-based estimation.)

The second way to overcome the restrictiveness of the IIA assumption is to use nested multinomial logit to model choice probabilities. The nested multinomial logit model is one in a class of generalized extreme value (GEV) models and captures correlations among error terms of different alternative-specific utilities by grouping alternatives with shared unobserved characteristics together into nests. These models have become common in the recreation demand literature. For some recent examples, see Parsons and Needelman (1992), Hausman, Leonard, and McFadden (1995); Needelman and Kealy (1995), and Parsons and Hauber (1998).

There are a wide variety of nesting structures one may choose in estimating a random utility model of recreation demand, yet no strict theoretical basis for choosing among them. As a result, nesting structures vary considerably across studies. Consider some examples. Hausman, Leonard, McFadden (1995) and Parsons and Needelman (1992) nest recreation sites geographically. McConnell and Strand (1994) nest recreation sites conditional upon the choice of fishing mode and target species. Parsons and Hauber (1998) do the reverse, modeling the choice of target species as conditional upon site choice. Similarly, Hausman, Leonard, McFadden (1995, p. 4) suggest modeling the choice of fishing mode as conditional upon site choice. And finally, numerous studies include a no-trip decision in the model in a nest separate from the basic site choice model (see Morey, 1998). This list is not exhaustive but demonstrates the variety of structures used.

Welfare analysis is the motivation for virtually all random utility models of recreation demand. Most applications involve valuing changes in catch rates of fish, changes in environmental quality, or loss of specific sites. It seems natural to ask, what effect does the choice of nesting structure have on the welfare estimates generated from these models? The research presented in this paper is an analysis of this effect. We use a data set of recreational fishing in Maine and compare the results of nine alternative nesting structures, including a non-nested specification. In all cases, we model an angler's choice among a common set of elemental alternatives. To date, only Kling and Thomson (1996) have analyzed explicitly the effect of nesting structure choice on welfare estimates. They find that different nesting structure specifications result in substantially different welfare estimates. It is useful to know whether this is a result that will persist with other data sets, nesting structures, and types of welfare analyses.

In the following section we present the assumptions underlying the specifications of the choice set employed in this study, as well as the functional form of the conditional indirect utility function. Also in Section 2, we describe the distributions of the stochastic error term used to generate the different nested models. The data are described in Section 3, and the estimation results are presented in Section 4. Also in Section 4, we evaluate the resulting parameter estimates against conditions for consistency with stochastic utility maximization. In Section 5, we examine the resulting welfare estimates and compare these results to the Kling and Thomson findings. We close with conclusions in Section 6.

2. Methodology

2.1 Specification of Deterministic Utility

We assume that each angler chooses a single element from a two-dimensional choice set composed of all combinations of site and target species available in the state of Maine. Therefore, the elemental alternative chosen by an individual is a site-species combination. The full choice set consists of 2029 fishing sites and up to four possible target species at each site (salmon, trout, bass, or perch)1; however, not all species are present at each site. The set of alternatives from which an individual chooses is $C = C_T - C_0 = 4629$; where C_T=8116 is the number of conceivable alternatives (2029x4); and C_0=3487 is the number of infeasible site-species combinations.2

In each model, an angler's conditional utility function for a fishing trip to site j targeting species f is assumed to be additively separable in the unobserved errors ϵ_{jf}, and takes the form

$$U_{jf} = \mu_{jf} + \epsilon_{jf}, \qquad (2.1.1)$$

where j = 1,...,J and f = s,t,b, or p (salmon, trout, bass, or perch). The term μ_{jf} is the deterministic portion of utility and is a function of explanatory variables and unknown parameters. μ_{jf} is unique to each angler;

1 The trout, bass, and perch targets are aggregates of more specific targets: Trout consists of brook trout, brown trout, lake trout, and other unspecified trout; Bass consists of largemouth bass, smallmouth bass, and other unspecified bass; and Perch consists of bullhead, muskie, pickerel, white perch, and other unspecified perch.

however, an individual's underlying preference ordering, captured by this deterministic portion of the indirect utility function, is the same for each of the nested models estimated in this study. μ_{jf} is assumed to be linear in the parameters and takes the form

$$\mu_{jf} = \alpha p_j + \beta \mathbf{x}_j + \gamma_f \mathbf{w}_{jf} + \delta \mathbf{z}_f, \qquad (2.1.2)$$

where α, β, γ_f, and δ are unknown parameters to be estimated and p_j, $\mathbf{x}_j$, $\mathbf{w}_{jf}$, and $\mathbf{z}_f$ are observed data.

The general form of the deterministic indirect utility function is an artifact of the assumptions underlying the choice set and the resulting elemental alternatives. In particular, because each alternative is assumed to be a site-species combination, some alternative-specific characteristics will vary only across sites, while others may vary among species as well. The term p_j is the sum of the angler's travel and time costs of reaching site j. The vectors $\mathbf{x}_j$ and $\mathbf{w}_{jf}$ are sets of site characteristics described in Table 1. The effect of $\mathbf{x}_j$ on utility is invariant with species choice while the effect of $\mathbf{w}_{jf}$ varies with both site choice and species choice. Finally, $\mathbf{z}_f$ is a vector intended to capture an angler's preference for specific species.

2.2 Nesting Structure Specifications: Distribution of the Error Term

The assumptions underlying the specification of the choice set and the resulting elemental alternatives also provide the basis for the development of the different nesting structures examined in this study. Some elements of the choice set are logically related and should be grouped together within a nest because they share a common element along one dimension of the choice set. For example, if it is assumed that all species at a given site share common unobserved characteristics, a nesting scheme in which species choice is conditional upon site choice should be used. This approach implies that if the species at a particular site have some set of unobserved characteristics in common, then we would expect that when an angler's demand for a particular site-species combination falls because some attribute of that particular alternative has been degraded, he will be more likely to increase his demand for targeting another species at the same site than to increase his or her demand for the same species at a different site. If instead it is assumed that all sites containing a particular species share common unobserved characteristics, then a

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2 1664 sites do not have salmon, 458 sites do not have trout, and 1365 sites do not have bass.

nesting specification in which site choice is conditional upon species choice should be used.

A second method of nesting arising from this choice set involves partitioning one dimension of the choice set into subsets. For example, the site dimension could be partitioned into geographic regions where it is assumed that sites within a particular region share certain unobserved attributes. In addition, the site dimension could also be partitioned into rivers and lakes and the species dimension could be partitioned in cold- and warm-water species. The former partitioning implies that anglers are likely to consider fishing trips to two different lakes to be better substitutes than one fishing trip to a lake and another to a river, while the latter suggests that anglers who target cold water species tend not to target warm water species and vice versa.

This study involves estimation of eight different nested discrete choice models and one non-nested model using a common data set and the common deterministic utility specification in equation (2.1.2).3 The two basic nested models appear in the tree diagrams in Figures 1 and 2. Figure 1 represents a two-level nesting structure in which site choice is modeled as conditional upon species choice. In this model, all sites at which a particular species is present are grouped together. This nesting structure (2-level species-site) implies the following cumulative joint distribution for the vector of errors $<\varepsilon_{jf}>$:4

$$F(<\epsilon_{jf}>) = \exp\{-\Sigma_{f=1}^{F}[\Sigma_{j=1}^{Jf}\exp(-\epsilon_{jf}/\rho_f)]^{\rho f}\} \qquad (2.2.1)$$

where ρ_f, $\forall f$=s,t,b,p, is the scale parameter on each species nest.

Figure 2 provides a representation of the same decision except that the nesting structure has been flipped; that is, the species choice is now conditional upon the site choice. While the structure presented in Figure 1 suggests that salmon fishing at site j is a better substitute for salmon fishing at site i than is bass fishing at site i, the nesting structure in Figure 2 suggests that bass fishing at site i is a better substitute for

3 While the number of possible nesting specifications is quite large, we choose to examine only nine in this study. Each of the nesting structures we examine is derived by exploiting the multidimensional nature of the choice set as described above; however, this set of nesting specifications is far from exhaustive.

4 Throughout this manuscript, the term 'species-site model' is intended to mean a model in which site choice is conditional upon species choice; that is, a model with species at the top of the nesting structure. The elemental alternative for a species-site model is labeled 'jf'. The reverse is true for the species-site model.

salmon fishing at site i than is salmon fishing at site j. This second nesting structure (2-level site-species) implies the following cumulative joint distribution for the vector of errors $< \varepsilon_{fj} >$:

$$F(<\epsilon_{fj}>) = \exp \{-\Sigma_{j=1}^{J} [\Sigma_{f=1}^{Fj} \exp(-\epsilon_{fj}/\rho_j)]^{\rho_j}\} \qquad (2.2.2)$$

in which ρ_j, $\forall j= 1,...,J$ is the scale parameter on each site nest.

The scale parameters capture the degree of correlation among alternatives within the nest such that the degree of correlation $= (1-\rho^2)$. In effect, the scale parameter provides an estimate of the degree of substitution among alternatives within the nest. The closer the scale parameter is to zero, the greater the substitution among alternatives within the nest. As the value of the scale parameter rises from zero within the unit interval, the degree of substitution among alternatives within the nest falls. When the scale parameter is unity, then the independence of irrelevant alternatives holds. In both equation 2.2.1 and 2.2.2, if the scale parameters are equal to unity ($\rho = 1$) then the nested model collapses to a non-nested multinomial logit model of the form

$$F(<\epsilon_{jf}>) = F(<\epsilon_{fj}>) = \exp \{-\Sigma_{f=1}^{F} \Sigma_{j=1}^{J} \exp(-\epsilon_{jf})\} \qquad (2.2.3)$$

The other models estimated in this study are developed as more general cases of the two described above. First, in both the species-site and site-species models, the species dimension of the choice set is partitioned such that the cold water species, salmon and trout, are grouped together. This grouping is made because we believe a priori that these species are close substitutes. As a result of this partitioning, we have two additional models, each with three levels of nesting: 3-level species-site (w/ cold) and 3-level site-species (w/ cold). The tree diagrams associated with these nesting specifications are presented in Figures 3 and 4.

We also partitioned the site dimension into rivers and lakes. This gives two more three-level nested models: 3-level species-site (w/ river&lake) and 3-level site-species (w/ river&lake). The tree diagrams associated with these nesting specifications are presented in Figures 5 and 6.

Finally, we estimate two four-level models which include both the cold water nest and the river and lake nests. These are our most elaborate models estimated in this study and are referred to as the 4-level species-site model and the 4-level site-species model. The tree diagrams for these models are presented in Figures 7 and 8.

It is common to restrict the scale parameters (ρ) to be constant across groupings within each level of the nest. For example, in the four-level model with site at the top (4-level site-species) it is assumed that $\rho_k = \rho_j = \rho_{site}$, where j indexes sites. Likewise, $\rho_r = \rho_l = \rho_{river/lake}$, and $\rho_s = \rho_t = \rho_b = \rho_p = \rho_{species}$. This convention is adopted throughout this study to avoid the computational burden that arises when the number of groups within each nesting level is large, as is the case in the 4-level model with site at the top in which each site is treated as a unique nest. 5

2.3 Nesting Structure Specification: Choice Probabilities

Given the specification of the deterministic portion of indirect utility and the specification of the distribution of the error term, the probability that an individual will choose species-site combination fj can be calculated. Using the two-level species-site nesting specification as an example, the probability that an angler will choose site j and target species f is

$$P(jf) = P(j|f) \bullet P(f). \qquad (2.3.1)$$

The probability of targeting species f at site j equals the probability of fishing at site j conditional upon the probability of targeting species f, multiplied by the probability of choosing species f. Given the distribution of the error term in equation (2.2.1) and deterministic utility specification described in equation (2.1.2), this choice probability is

$$P(j|f) = \exp\,(\alpha p_j + \beta x_j + \gamma_f w_{jf})/\rho_f) \;/\; I_f$$

$$P(f) = \exp\,(\delta z_f + \rho_f \ln(I_f)) / \,[\Sigma_{g=1}{}^{F} \exp\,(\delta z_g + \rho_g \ln(I_g))]$$

where I_g takes the form

$$I_g = \Sigma_{i=1}{}^{Ng} \exp\,((\alpha p_i + \beta \mathbf{x}_i + \gamma_f w_{ig})/\rho_g. \qquad (2.3.2)$$

I_f is the value of I_g for the case g = f and $\Sigma_{i=1}{}^{Ng}$ sums over N_g sites in the species g nest.

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5 Parsons and Hauber (1998) estimate a three-level model with site at the top and allow the scale parameter to vary between rivers and lakes at the site level.

Likewise, uisng the 2-level site-species model in equation (2.2.2) and calculate the choice probability as the probability of targeting species f, conditional on choosing site j, multiplied by the probability of choosing site j; we have,

$$P(fj) = P(f|j) \bullet P(j)$$

where

$$P(f|j) = \exp((\gamma_f w_{fj} + \delta z_f)/\rho_j) / Ij,$$

$$P(j) = \exp(\alpha p_j + \beta x_j + \rho_j \ln(I_j)) / [\Sigma_{i=1}^{N} \exp(\alpha p_i + \beta x_i + \rho_i \ln(I_i))], \text{ and}$$

$$I_i = \Sigma_{g=1}^{F} \exp((\gamma_g w_{gi} + \delta z_g)/\rho_i). \quad (2.3.3)$$

I_j is the value of I_i for the case i = j and $\Sigma_{g=1}^{F}$ sums over F species in the site nest.

All parameters are estimated by full information maximum likelihood. A unique vector of coefficient estimates is derived for each error specification. For each model, the corresponding coefficient estimates can then be used to estimate deterministic utility under different states of the world (i.e., before and after a change in site quality) and the compensating variation associated with such a change as described below. Because the number of sites is large, we estimate each of the seven models using a randomly drawn subset of sites (see Parsons and Kealy, 1992). A total of 30 sites is used in estimation: the actually chosen site plus 29 sites drawn from all sites in the choice set.6 Welfare estimation is always done over the full set of sites using the parameter estimates from the random draws. Each model is estimated twenty times using different random draws (see Feather, 1994; and Waters and Dietz, 1995).

3. Data

The data set consists of 2425 fresh-water fishing day trips taken by 143 Maine anglers during the summer and fall of 1989. These data come from a survey conducted in 1989 for the National Acid Precipitation Assessment Program (NAPAP). The details of the survey are presented in Shankle et al.

6 In the species-site models we included the weighting term ln(N/M) in each site utility, where N is the full number of sites in the nest set and M is the number of randomly drawn sites in the nest. This adjustment accounts for randomly drawing sites. However, the differences between models using and not using this weighting term were negligible. Hence, we report parameter estimates from the randomly drawn

(1990). For each individual, data are available on the number of trips taken, the site visited on each trip, and the species of fish targeted on each trip. Demographic data, including income and information regarding each angler's favorite target species, are also included.

Characteristic data are available for 1899 lakes in Maine and include information on water quality and the abundance of each species, as well as depth, elevation and location. Lake data come from the *Maine Lake Survey* prepared by the Maine Department of Inland Fisheries and Wildlife. Approximately 98% of the lakes visited in the NAPAP survey are covered in the lake survey.

Characteristic data are also available for 130 rivers reaches in Maine. Larger rivers are divided into smaller reach segments to capture differences in location and site characteristics. The river data include information on location, the presence of salmon, and a dummy variable indicating whether the site is a reach on one of the major rivers in the state. The data set for river sites is somewhat limited because it was compiled using several Maine fishing guides.

Water quality data were obtained from the Maine 305(b) report to the US Environmental Protection Agency (EPA) for 1990: *State of Maine, 1990 Water Quality Assessment.* In this report, the state identifies lakes that fail to attain certain water quality standards, primarily due to non-point source pollution. The state also identifies rivers with elevated levels of toxics.

Finally, to estimate these models, travel times and distances for each angler to all sites in the choice set are necessary. These values were determined using the software package HYWAYS/BYWAYS.

4. Estimation Results

4.1 Coefficient Estimates

The coefficient estimates, while not the primary focus of this research, are indicative of the applicability of these models to the choices of site and species made by anglers in the state of Maine. Throughout each of the nine models and across all twenty draws, the results of the parameter estimation are generally consistent with expectations. The coefficient on the price term is always negative and significant,

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choice set without weighting. The welfare measures are, as usual, derived over the full choice set.

indicating that the likelihood of traveling to a particular site for the purpose of recreation is inversely related to the cost of reaching the site. Not surprisingly, the coefficients on the species preference variables are always positive and significant. The parameters on the lake-specific variables, ln($ACRES_j$), BOAT*$ACCESS_j$, and $REMOTE_j$ are always positive and significant, as are those on the variables representing salmon, trout and bass abundance, $ABUND_{jf}$, f=s,t,b. The coefficients on AGE*$REMOTE_j$, and $DIRT_{jf}$ are consistently negative and significant. Likewise the parameter estimates for the river-specific variables $MAJOR_j$ and R_j are positive and significant.

There are two instances, however, in which the estimation results differ from expectations. The coefficient on the lake-specific variable for perch abundance, $ABUND_{jf}$, always negative and significant. At first this appears surprising as it was expected that an increase in the availability of a species would increase the utility associated with a particular site. However, upon further examination, this result may make sense. Since perch abundance is a dummy variable equal to one for a large group of warm water species other than bass, it may be the case that anglers targeting one species in the group care little for, and perhaps even dislike, the other species included in the group.

Most troublesome is the estimated coefficient on the variable R_j * $TOXIC_j$, as it is a critical element in the welfare evaluation of a river clean-up. This parameter estimate is expected to be negative as a toxic river site should yield less utility than a clean river site. However, the parameter estimate varies from positive to negative across draws and is often insignificant. This finding could be an artifact of the construction of the toxic variable. A river is labeled toxic if there is a fish consumption advisory associated with the site. It is assumed that an angler's perception of the site characteristics determines the effect a site's attributes have on his or her indirect utility. Because toxics may not be easily perceived, or if the fish consumption advisory is not widely disseminated, individual anglers may be unaware that a particular site is considered toxic. If so, it is expected that the toxic designation would have little effect on the utility associated with the site.

The parameters of primary interest in this study are the inclusive value (scale) parameters. As mentioned above, the scale parameters capture the degree of correlation among alternatives within a nest.

McFadden (1981) has shown that $0 < \rho \leq 1$ is a globally sufficient condition for the nested model to be consistent with utility maximization. Borsch-Supan (1990) identifies local sufficiency conditions permitting $\rho > 1$, and Kling and Herriges (1995) and Herriges and Kling (1996) extend the work of Borsch-Supan by developing and implementing empirical tests for the locally sufficient conditions for the consistency of nested models with utility maximization. The result of these efforts has been to expand the bounds of acceptable values of ρ, but only slightly. In particular, Herriges and Kling find that $1 < \rho < 2$ will more often than not fail to meet the Borsch-Supan conditions.7

Atherton et al. (1990) and Hausman, Leonard, and McFadden (1995) provide an alternative interpretation of $\rho > 1$. From a purely statistical perspective, if $\rho > 1$ then there is greater correlation among the utilities of elements of different nests than among the utilities of elements within the same nest. For example, in the two-level model with site at the top, an estimated scale parameter on the site nest, ρ_{site}, greater than one would indicate that a given species at site i is a better substitute for the same species at site j than is another species at site j. While estimates of $\rho > 1$ may not be wholly consistent with stochastic utility maximization per se, models in which scale parameters exceed one still capture the nature of substitutability among alternatives and may thereby represent individual choice behavior.

The estimates of the scale parameters in this study are consistently positive and statistically significantly greater than zero at a 1% level of confidence. However, these estimates do not always lie within the unit interval. Table 2 summarizes the scale parameter estimates relative to the conditions described above. Columns (3) - (5) display the minimum, median, and maximum values of the scale parameter estimates for each model across the twenty draws. Columns (6) - (8) show the number of times in twenty draws that an estimate is statistically significantly greater than one.

One conclusion that can be drawn from the results in column (6) is that in the models in which site choice is conditional upon species choice (species-site), the scale parameter on the species nests, $\rho_{species}$, lies within the unit interval in a large majority of the draws. In contrast, in the models in which species

7 We conducted the Kling and Herriges tests on a number of our two-level models and found no instance in which the result changed the conclusions presented later in this section.

choice is conditional upon site choice (site-species), the scale parameter on the site nests, ρ_{site}, is more often than not greater than one. If we use the McFadden (1981) criterion to evaluate these results, we conclude with our data that grouping alternatives sharing a common species into a nest is generally consistent with utility maximization while nesting by grouping alternatives sharing the same site tends not to be.

The scale parameter estimates on the higher order nests provide additional information as to the consistency of the nesting structures with utility maximization. As shown in column (7), in a large number cases in both the species-site and site-species models the scale parameter on the cold water nest, ρ_{cold}, exceeds one with statistical significance. These results suggest that bass and perch are closer fishing substitutes for salmon and trout than we had expected and that the cold water nest may be inappropriate.

Inclusion of the river-lake partition also gives interesting results. The scale parameter on the river-lake nes, $\rho_{river/lake}$, in the species-site models is usually positive and less than one (see column (8)), indicating that there is greater substitution among rivers or among lakes than between the two. In these models the river/lake split is consistent with utility theory. However, this does not hold in the site-species model in which $\rho_{river/lake}$ usually statistically significantly greater than one, indicating that there is greater substitution between rivers and lakes than there is among rivers and among lakes. This latter result may be capturing some of the substitution across sites for the same species.

Using consistency with utility theory as a guide then, we find that the preferred nesting structure is the 3-level species-site model that includes the river and lake nests. Among all the models, this is the only one that consistently resulted in scale parameter estimates within the acceptable range. It is also important to note that this model satisfies the less well know condition that the size of the ρ's decline as you move from the bottom to the top of the nest. Several of the other models violated this condition. See Ben-Akiva and Lerman (1985, p. 293) for more on this condition.

The results in Table 2 also reveal, however, that even though there are many cases in which the median estimated scale parameters across the twenty draws is greater than one, these estimates are usually close to one. For example, the highest median estimate of a scale parameter across all eight models is 1.27. This result differs significantly from that of Kling and Thomson who find that the scale parameters often exceed two and are sometimes greater than four. The implications of this result will be discussed in the next section.

4.2 Welfare Estimates

The welfare effect of a change in site quality can be represented as the difference between the maximum expected utility of a trip, with and without the change, divided by the estimated coefficient on the travel cost variable (Hanemann, 1982; Morey, 1998). If it is assumed that income effects are zero, then the closed-form expression for compensating variation associated with a change in the quality of one or more sites, or the elimination of a site, takes the form

$$CV = EV = [EU^1 - EU^0]/-\alpha, \qquad (4.2.1)$$

where α is the coefficient on the travel cost term (see Table 1) and can be interpreted, in absolute value, as the marginal utility of income. EU^0 and EU^1 represent the expected utility of a trip without and with the change, respectively.

In a simple two-level model in which site choice is conditional upon species choice, EU^a takes the form

$$EU^a = \ln[\Sigma_{f=1}{}^F \exp(\delta z_f + \rho_{species} \ln(I_f))] + .57 \qquad (4.2.2)$$

where $I_f = \Sigma_{i=1}{}^{Nf} \exp((\alpha p_i + \beta x_i + \gamma_f w_{if})/\rho_{species})$ is the inclusive value for species nest f, a is equal to 0 or 1, and $\rho_{species}$ is the scale parameter on the species nests, capturing the degree of substitution among sites within each species nest.

For each of the nine models, welfare estimates are derived for three common policy scenarios. These policy scenarios include the clean-up of all lakes that do not meet EPA quality standards, the clean up of all toxic river sites, and the elimination of salmon as a possible target species. The first two scenarios each provide an estimate of the benefits of a hypothetical pollution clean-up or abatement program, while the third provides an estimate of the overall value of salmon fishing day trips. In the lake clean-up scenario, all polluted lakes are assumed to be cleaned. Because cold water species (salmon and trout) are often greatly affected by non-point source pollution while other species such as bass and perch remain unaffected, the lake clean-up has the added effect of making both salmon and trout abundant in the formerly dirty lakes even if neither of these species was present before the clean-up. In the river clean-up, all rivers are assumed to be free of toxins, thereby removing all fish consumption advisories. Salmon elimination is accomplished by removing salmon from all sites at which it was formerly present. The median per trip values across the twenty random draws for each model for all three welfare scenarios are presented in Table 3.

The estimates of per trip value for cleaning polluted lakes range from \$3.11 per person in the four-level site-species model to \$1.89 in the non-nested model. The value for cleaning up toxic river sites ranges from \$1.25 per person in the two level site-species and three level site-species (cold) model to \$1.14 in the non-nested model. The estimates of welfare loss associated with the elimination of salmon fishing range from \$4.01 per person in the four-level species-site model to \$2.74 in the non-nested model.

The variation in welfare estimates across the models is not large. Among the nested models only, the difference between the highest and lowest median estimates is 9.31% in the lake clean-up scenario, 8.08% in the river clean-up scenario, and 37.28% in the salmon elimination scenario. These findings run counter to Kling and Thomson (1996) who found that welfare estimates can vary widely across alternative nested models by up to 174%.8 The difference between our results and those of Kling and Thomson appear to originate with the estimates of the scale parameters. We find that the scale parameters are often close to one, and that there is not a great deal of variation across the models. Kling and Thomson find much more variation in the scale parameters across their models. In particular, their parameters that exceed one are often much greater that one. In a couple cases the parameters are greater than four. Our interpretation of the difference in these findings is as follows.

When nesting structures are specified correctly (consistent with utility theory), substitution patterns are revealed through scale parameters lying within the unit interval. The closer the scale parameter is to zero, the greater the degree of substitution among sites within the nest. When nesting structures are

8 Note that we could have used the Krinsky-Robb technique to estimate the standard error of the welfare estimates (see Krinsky and Robb, 1986, 1991; Kling and Thomson, 1996; and Parsons and Kealy, 1994) but we cannot compare the differences between welfare estimates in formal statistical tests because they are not

specified incorrectly, the *same* substitution patterns are revealed through scale parameters greater than one. The high scale parameter estimates in the incorrect model are signaling that there are poor substitutes within each nest, and revealing that there is a great deal of cross-nest substitution present in the data. The greater the size of the scale parameter, the greater is the degree of cross-nest substitution.

It stands to reason that data sets having strong correlation patterns among the random utilities will tend to have highly variable scale parameter estimates as the nesting structures are altered. These estimates will be low (less than one) when the nests are specified accurately and high (much greater than one) when specified inaccurately. This point is confirmed by Herriges and Kling (1998) in a rather novel Monte Carlo study. They used fabricated data for which they know the parameters of site utility, the correct nesting structure, and size of the scale parameters in the nests. When a model is estimated using the fabricated data but an incorrect nesting structure, the scale parameter is larger than one, signaling the cross-nest substitution which we know is true. The lower the known scale parameter is set in the true model, the higher the estimated scale parameter is in the misspecified model.

It appears as though the Kling and Thomson data have some strong correlation patterns along the lines they are considering for nesting. This creates the wide variability in their scale parameter estimates as the nesting structures are altered. The large varation in the scale parameter estimates translate into large variation in the welfare estimates across the models. When correlation patterns in the data are strong, accurately accounting for them is important for welfare estimation. Again, this result is confirmed by Kling and Herriges (1998) in their Monte Carlo study. When their true models had low scale parameters, their misspecified models not only had large scale parameters, but also reported welfare estimates which deviated widely from the true values. In our analysis, in contrast with Kling and Thomson, the scale parameters tend to be closer to one. The actual correlation patterns in our data set appear to be less important. Hence, when we misspecify our model the scale parameters are larger than one, but not to an extent realized by Kling and Thomson. Since correlation is less important in our data set (at least along the lines we a considering), accounting for it accurately is less important for the welfare estimates. Indeed,

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independent random variables. Kling and Thomson attempted no such formal test either.

there is little loss in accuracy across our models from using any of the nested models. 9

It is important to note that correlation among alternatives within a nest may result from measurement inaccuracies as well. If the researcher simply lacks good information on rivers for example, it is more likely that rivers will have shared unobserved characteristics leading to scale parameter estimates that will be sensitive to the position of rivers in the nesting structure. Hence, large variation in scale parameter estimates and, in turn, welfare estimates is likely to be more common in models with simple specifications missing attributes which matter to individuals.

It is also worth noting that the welfare estimates from the nested logit models in our analyses are larger than the estimates from the non-nested models. This appears to be due to the welfare scenarios we are considering. All three cases consider large changes within specific nests which generate substitution across nests. Since the nested models are accounting for substitution within the nests, this cross nest substitution generates larger values in the nested models. Small changes at one or two sites within a nest would have produced the opposite effect. Again, the differences here are not dramatic.

In addition, Kling and Thomson present a series of likelihood ratio tests comparing their different nesting structures. They use a test suggested by Pollock and Wales (1991). They find that the models which are least consistent with utility theory dominate (in a statistical sense) the models which are consistent with utility theory. That is, the models with largest scale parameters greater than one tended to have the highest likelihood values. This finding is consistent with that of Ben-Akiva and Lerman (1985, pp 317-19) and we find the same in our models. In our case, the species-site models are 'most' consistent with utility theory, but the site-species models dominate in statistical tests. However, we are skeptical of these tests. They assume that the likeihood values from each model are based underlying probabilities generated from a well-behaved multivariate pdf. All of the models with scale parameters falling outside the unit interval do not, by definition, have well-behaved distributions since they permit negative and

9 It is possible that other nesting structures using our data would have generated scale parameters which deviated further from 1. For example, if we had considered nesting by geographic regions or modes, we may have found scale parameters much closer to 0 or significantly larger than 1.

greater than zero probabilities. This being the case, we cannot be certain that the likelihood ratios have a Chi-squared distribution nor that they are valid for use in statistical tests.

5. Conclusions

In Kling and Thomson's closing comments they remark that

> "Many of the results in this data set may not be generalizable to other recreation demand applications. To this end, it would be useful to examine the sensitivity of welfare measures for other data sets to the same of set of specification issues here. Only by examining the magnitudes of specification differences across a variety of data sets can any generalizations be fruitfully made."

While they consider a number of different specification issues with nested logit models, we focus our analysis on the sensitivity of welfare estimates to changes in nesting structure. Our results suggest that their finding that welfare estimates are highly sensitive to choice of nesting structure is not a general result. In our application the models are not particularly sensitive to the change in structure.

Taken together the two studies are rather instructive. We have argued that the variation in the welfare estimates in both studies largely tracts the variation in the scale parameter estimates across the models. Kling and Thomson found wide variation in their scale parameter estimates and we found relatively little variation. We argue that the greater the degree of actual correlation among random utilities, the greater is the likelihood in observing wide variation in scale parameter estimates and, in turn, wide variation in welfare estimates across different nesting structures. In a data set where strong patterns of correlation (or substitution) exist and the researcher misspecifies the nesting structure, scale parameter estimates are likely to exceed 1 by a wide margin.

In such models, the welfare estimates would appear to be on less firm ground. The high scale parameter estimate signals that there is an alternative nesting structure which accounts more accurately for the correlation that exists in the data *and* that the misspecified model is likely to report estimates which deviate widely from the prefer structured. This result is confirmed in a Monte Carlo study by Herriges and Kling (1998).

The implication is one that many analysts may have suspected. Nested logit models with scale parameter estimates in excess of 1, but only slightly so, perhaps less than 1.5, are not likely to give terribly biased welfare estimates. Models with large scale parameter estimates, on the other hand, may give extremely biased welfare estimates and caution is warranted.

The discussion here is not intended to imply that analysts should conduct a regression 'hunt' for low scale parameter estimates in every instance or even that having a model with low scale parameters is evidence of having found the true model (no doubt other structures with scale parameters lying on the unit interval may still exist). Rather it is to provide some evidence for understanding and interpreting welfare results from nested logit models.

REFERENCES

Adamowicz, W.L. 1994. "Habit Formation and Variety Seeking in a Discrete Choice Model of Recreation Demand, " *Journal of Agricultural and Resource Economics* 19: 19-31.

Atherton, T., M. Ben-Akiva, D. McFadden, and K. Train. 1990. "Micro-Simulation of Local Residential Telephone Demand Under Alternative Service Options and Rate Structures," in A. Fontenay *et al.*, eds. *Telecommunications Demand Modelling: An Integrated View*. New York: North Holland.

Bockstael, N.E., W.M. Hanemann, and C.L. Kling. 1987. "Estimating the Value of Water Quality Improvements in a Recreation Demand Framework, " *Water Resources Research* 23: 951-960.

Borsch-Supan, A. 1990. "On the comparability of Nested Logit Models With Utility Maximization, " *Journal of Econometrics* 43: 373-388.

Ben-Akiva, M. and S.R. Lerman. 1985. *Discrete Choice Analysis: Theory and Application to Travel Demand*. MIT Press, Cambridge.

Chen, H.Z. and S.R. Coslett. 1998. "Angler's Perceptions of Environmental Quality and Benefit Estimation in Multinomial Probit Models: A simulation Approach, " *American Journal of Agricultural Economics* forthcoming.

Feather, P.M. 1994. "Sampling and Aggregation Issues in Random Utility Model Estimation, " *American Journal of Agricultural Economics* 76: 722-779.

Hanemann, W.M. 1982. "Applied Welfare Analysis with Qualitative Response Models, "California Agricultural Experiment Station.

Hanemenn, W.M. 1978. "A Methodological and Empirical Study of the Recreation Benefits from Water Quality Improvements, " Ph.D. dissertation, Department of Economics, Harvard University.

Hausman, J.A., G.K. Leonard, and D. McFadden. 1995. "A Utility Consistent, Combined Discrete-Choice and Count Data Model Assessing Recreational Use Losses Due to Natural Resource Damage," *Journal of Public Economics* 56: 1-30.

Herriges, J.A., and C.L. Kling. 1996. "Testing the Consistency of Nested Logit Models with Utility Maximization" manuscript.

Herriges, J.A., and C.L. Kling. 1998. "The Performance of Nested Logit Models When Welfare Estimation is the Goal," *American Journal of Agricultural Economics, 79: 792-802.*

Jones, C.A. and Y.D. Sung. 1991. "Valuation of Environmental Quality at Michigan Recreational Fishing Sites: Methodological Issues and Policy Applications." University of Michigan.

Kling, C.L., and J.A. Herriges. 1995. "An Empirical Investigation of the Consistency of Nested Logit Models with Utility Maximization," *American Journal of Agricultural Economics* 77: 875-884.

Kling, C.L., and C.J. Thomson. 1996. "The Implications of Model Specification for Welfare Estimation in Nested Logit Models," *American Journal of Agricultural Economics* 78: 103-114.

Krinsky, I., and A. Robb. 1986 "On Approximating the Statistical Properties of Elasticities," *Review of Economics and Statistics* 68:715-719 (correction, 1990 *Review of Economics and Statistics* 72: 189-190).

McConnell, K.E., and I.E. Strand. 1994. *The Economic Value of Mid and South Atlantic Sportfishing. Volume II.* Report on EPA Cooperative Agreement, University of Maryland.

McFadden, D. 1981. "Econometric Models of Probabilistic Choice, " in C.F. Manski and D. McFadden eds. *Structural Analysis of Discrete Data.* Cambridge: MIT Press.

McFadden, D. 1978. "Modelling the Choice of Residential Location, " in A. Karquist *et al.*, eds., *Spatial Interaction Theory and Planning Models*. Amsterdam: North-Holland.

Morey, E.R. 1998. "Two RUMs UnCLOAKED: Nested-Logit Models of Site Choice and Nested-Logit Models of Participation and Site Choice." forthcoming Kling and Herriges book on recreation demand.

Morey, E.R., R.D. Rowe, and M. Watson. 1993. "A Repeated Discrete Choice Model of Atlantic Salmon Fishing," *American Journal of Agricultural Economics.* 75: 578-592.

Needelman, M.S., and M.J. Kealy. 1992. "Recreational Swimming Benefits of New Hampshire Lake Water Quality Policies: An Application of a Repeated Discrete Choice Model," *Agricultural and Resource Economics Review* 24: 78-87.

Parsons, G.R., and A.B. Hauber. 1998 "Spatial Boundaries and Choice Set Definition in Random Utility Models of Recreation Demand," *Land Economics*.

Parsons, G.R., amd M.J. Kealy. 1994. "Benefits Transfer in a Random Utility Model of Recreation, " *Water Resources Research* 30(8): 2477-2484.

Parsons, G.R., and M.J. Kealy. 1992. "Randomly Drawn Opportunity Sets in a Random Utility Model of Lake Recreation, " *Land Economics* 68 (1992): 93-106.

Pollack, R., and T. Wales. "The Likelihood Dominance Criterion," *Journal of Econometrics* 47: 227-242.

Shankle, S.A., J.E. Englin, W.H. Monroe, D.M. Beck, M.F. Casinelli, H. Saunders, and S.E. Grover. 1990. *Freshwater-Based Recreation in the Upper Northeast: Data Collection and Documentation.* Pacific Northwest Laboratory, Batelle Memorial Institute.

Stern, Steven. 1997. "Simulation-Based Estimation," *Journal of Economic Literature*, 35 (December 1997), 2006-2039.

Waters, S.M., and K.J. Dietz. 1995. "Measuring the Stability of Choice Set Randomization in Discrete Choice Models," manuscript, Triangle Economic Research.

TABLE 1

Functional From for Deterministic Utility (μ_{jf})

$$\mu_{jf} = \alpha p_j + \beta x_j + \gamma_f w_{jf} + \delta z_f$$

Variables Affecting Site Choice (p_j, x_j)

$$\alpha p_j = \alpha \, PRICE_j$$

$PRICE_j$	=	Opportunity Cost of Time plus Travel Cost of Reaching Site j {Hourly Wage * (Round Trip Travel Time in Hours) + (.30 * Round Trip Distance in Miles)}
		Note on Hourly Wage: For people on fixed incomes, Hourly Wage = (Annual Income)/2080. For Homemakers, retired, unemployed, or students, we assume Hourly Wage = $10.

$$\beta x_j = \beta_1 L_j*\ln(ACRES_j) + \beta_2 L_j*REMOTE_j + \beta_3 L_j*AGE*REMOTE_j + \beta_4 L_j*BOAT*ACCESS_j + \beta_5 R_j + \beta_6 R_j * MAJOR_j + \beta_7 R_j*TOXIC_j$$

$ACRES_j$	= Surface area of lake j in acres
$REMOTE_j$	= 1 if site j is reached only by off-road vehicle or on foot; 0 otherwise
AGE	= Age of respondent in years
BOAT	= 1 if individual fished from a boat on reported trip; 0 otherwise
$ACCESS_j$	= 1 if outboard motors are prohibited or restricted at site j; 0 otherwise
$MAJOR_j$	= 1 if river reach is on one of the major rivers in the state; 0 otherwise
	Major rivers include the Allagash, Androscoggin, Aroostook, Dead, Kennebec, Mattawamkeag, Moose, Penobscot, Saco, St. John, Sandy, Presumpscot, Sebasticook, Union, Piscataquis, and St. Croix.
$TOXIC_j$	= 1 if river reach j has elevated levels of toxics; 0 otherwise
R_j	= 1 if site j is a river or stream; 0 otherwise
L_j	= 1 if site j is a lake; 0 otherwise

TABLE 1

Functional From for Deterministic Utility (μ_{jf}), continued

Variables Affecting Species Choice (w_{jf}, z_f)

$$\gamma_f w_{jf} + \delta z_f = \gamma_{1f} L_j{*}DIRT_{jf} + \gamma_{2f} L_j{*}ABUND_{jf} + \delta_1 FAV_f + \delta_2 BEST_f$$

$DIRT_{jf}$	= 1 if site j is in nonattainment of EPA standards; 0 otherwise (bass and perch are believed to be unaffected by nonpoint source pollution in lakes)
$ABUND_{jf}$	= 1 is site j has species f in abundance; 0 otherwise
FAV_f	= 1 is species f is among the angler's 3 favorite species to target; 0 otherwise
$BEST_f$	= 1 if species f is the angler's favorite species to target; 0 otherwise

TABLE 2: Estimates of the Scale Parameters

(1) Species-Site Models	(2) Scale Parameter	(3) Minimum	(4) Median	(5) Maximum	(6) Number of Times in 20 Draws $\rho_{species} > 1^*$	(7) Number of Times in 20 Draws $\rho_{cold} > 1^*$	(8) Number of Times in 20 Draws $\rho_{river/lake} > 1^*$
2-Level Species-Site	$\rho_{species}$	0.5464	0.7539	1.2236	2		
3-Level Species-Site	$\rho_{species}$	0.5601	0.8406	1.2381	3		
(cold water nest)	ρ_{cold}	0.9557	1.0629	1.2791		17	
3-Level Species-Site	$\rho_{species}$	0.4286	0.7369	1.5170	1		
(river & lake nests)	$\rho_{river/lake}$	0.5369	0.7518	1.1462			1
4-Level Species-Site	$\rho_{species}$	0.4350	0.7435	1.5357	3		
	ρ_{cold}	0.9284	1.0401	1.2984		12	
	$\rho_{river/lake}$	0.5556	0.8333	1.1840			2

(1) Site-Species Models	(2) Scale Parameter	(3) Minimum	(4) Median	(5) Maximum	(6) Number of Times in 20 Draws $\rho_{site} > 1^*$	(7) Number of Times in 20 Draws $\rho_{cold} > 1^*$	(8) Number of Times in 20 Draws $\rho_{river/lake} > 1^*$
2-Level Site-Species	ρ_{site}	0.4746	1.1646	1.8968	13		
3-Level Site-Species	ρ_{site}	0.5163	1.2668	2.0029	14		
(cold water nest)	ρ_{cold}	0.4668	1.0409	1.6775		7	
3-Level Site-Species	ρ_{site}	0.7006	1.2736	2.1211	16		
(river & lake nests)	$\rho_{river/lake}$	0.7439	1.1509	1.5726			12
4-Level Site-Species	ρ_{site}	0.6887	1.1947	2.1913	13		
	ρ_{cold}	0.5584	1.0891	1.8364		10	
	$\rho_{river/lake}$	0.6353	1.1252	1.5551			13

TABLE 3: Median Per-Trip Welfare Estimates*

Model	Scenario 1	Scenario 2	Scenario 3
Non-nested	$1.89	$1.14	$2.74
2 Species-Site	$2.62	$1.17	$3.68
3 Species-Site (Cold)	$2.62	$1.16	$3.90
3 Species-Site (R/L)	$2.63	$1.19	$3.68
4 Species-Site	$2.58	$1.21	$4.01
2 Site-Species	$2.58	$1.25	$3.17
3 Site-Species (Cold)	$2.96	$1.25	$2.75
3 Site-Species (R/L)	$3.10	$1.16	$3.42
4 Site-Species	$3.11	$1.17	$3.00

* For each model we have 20 estimates corresponding to our 20 random draws. We calculated the average per trip value across all individuals in the sample for each model and here we report the median value across the 20 models.

Notes:
Scenario 1: Clean-up all dirty lakes and restore salmon and trout to abundance.
Scenario 2: Clean-up all toxic rivers having fish consumption advisories.
Scenario 3: Eliminate salmon fishing at all sites.

FIGURE 1

2-Level Model (species-site)

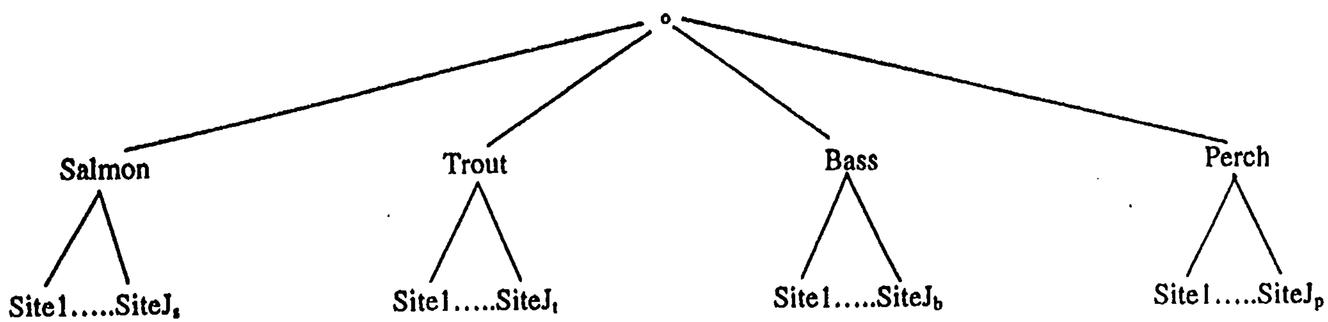


FIGURE 2

2-Level Model (site-species)

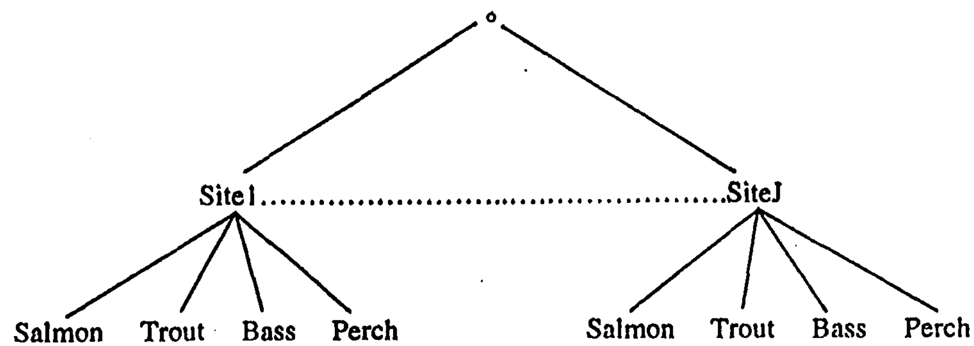


FIGURE 3

3-Level Model (species-site)
(with cold water nest)

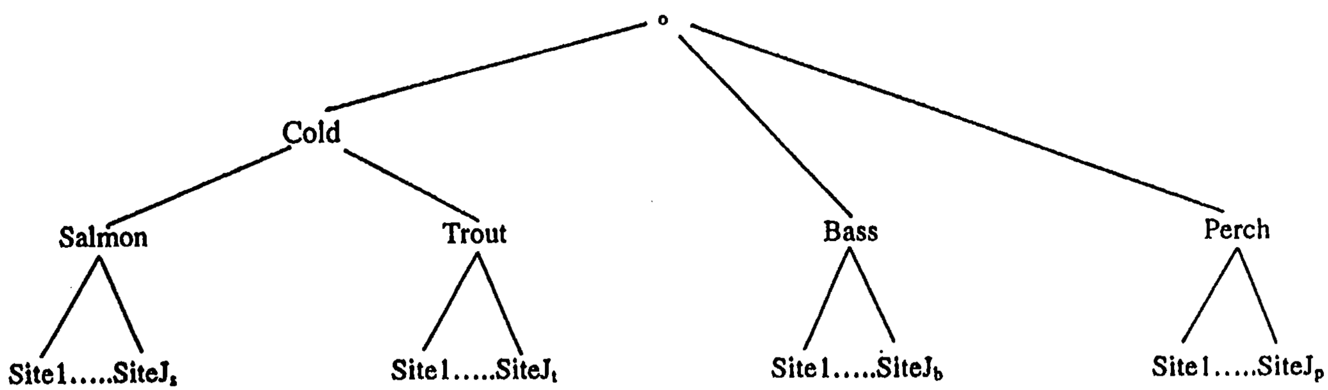


FIGURE 4

3-Level Model (site-species)
(with cold water nest)

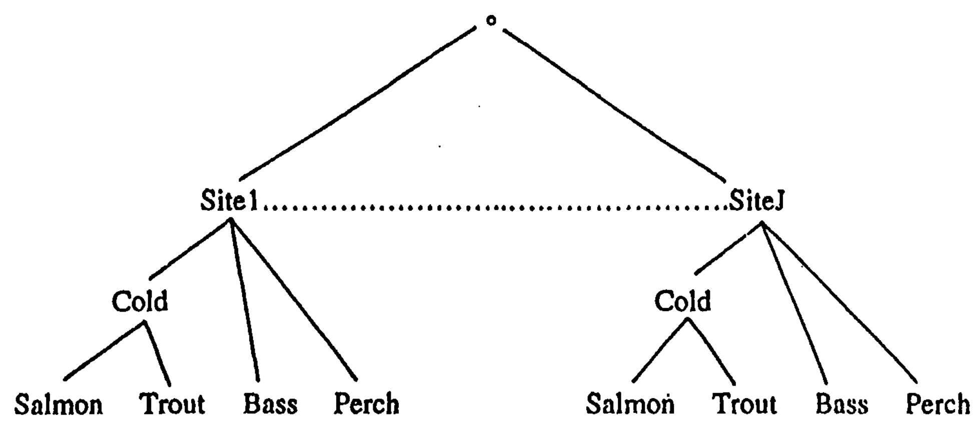


FIGURE 5

3-Level Model (species-site)
(with river and lake nests)

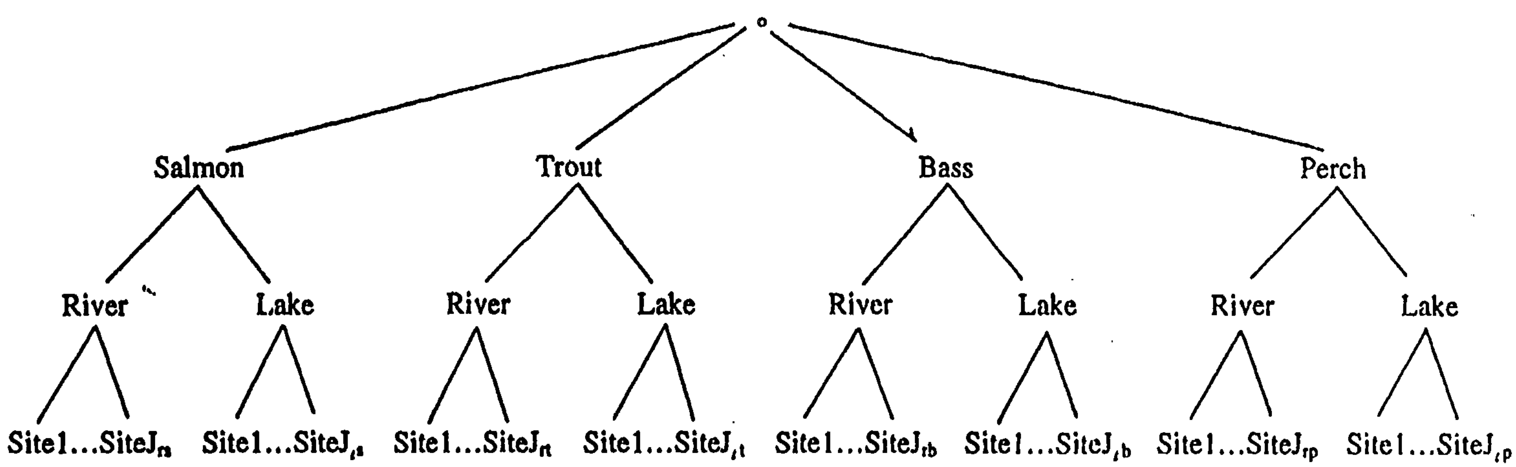


FIGURE 6

3-Level Model (site-species)
(with river and lake nests)

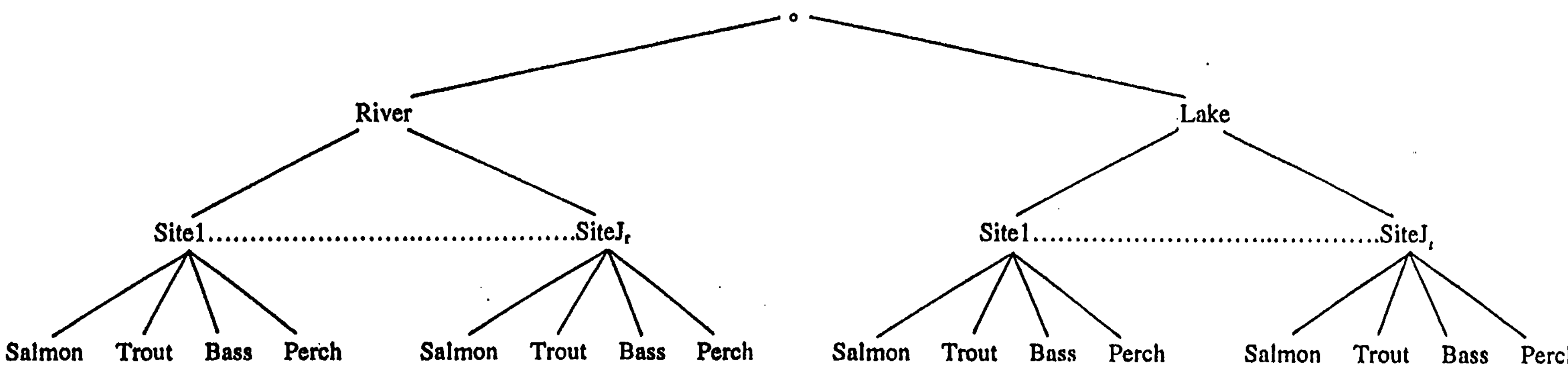


FIGURE 7

4-Level Model (species-site)

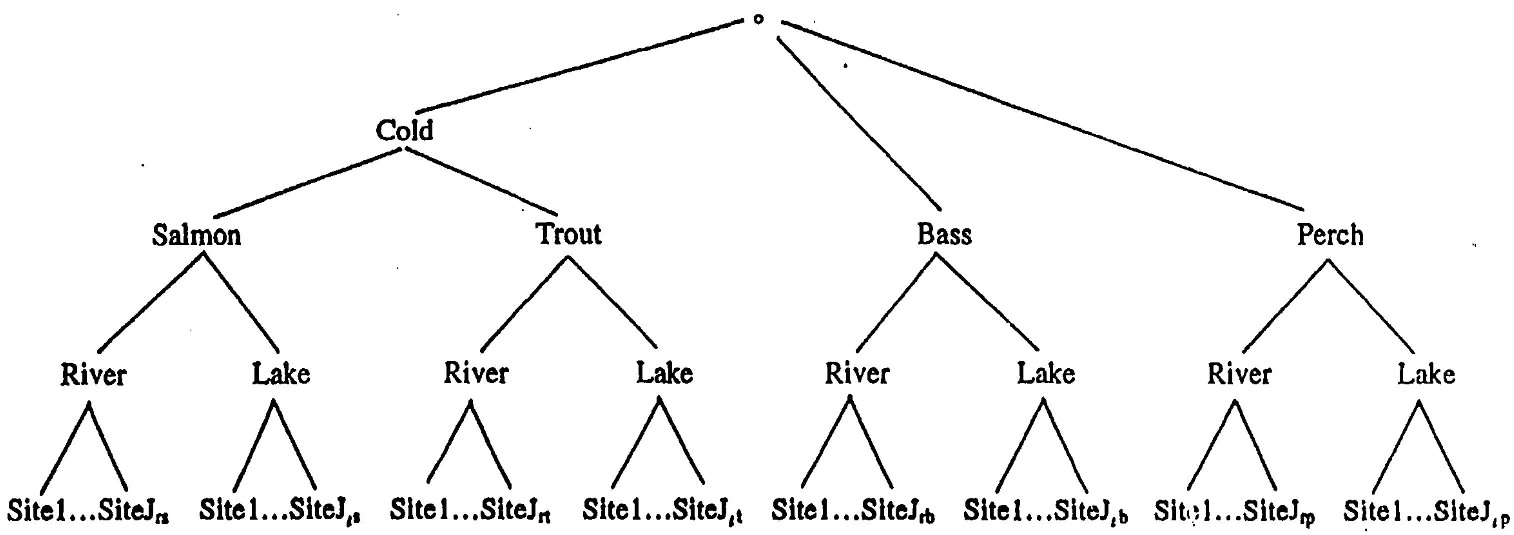
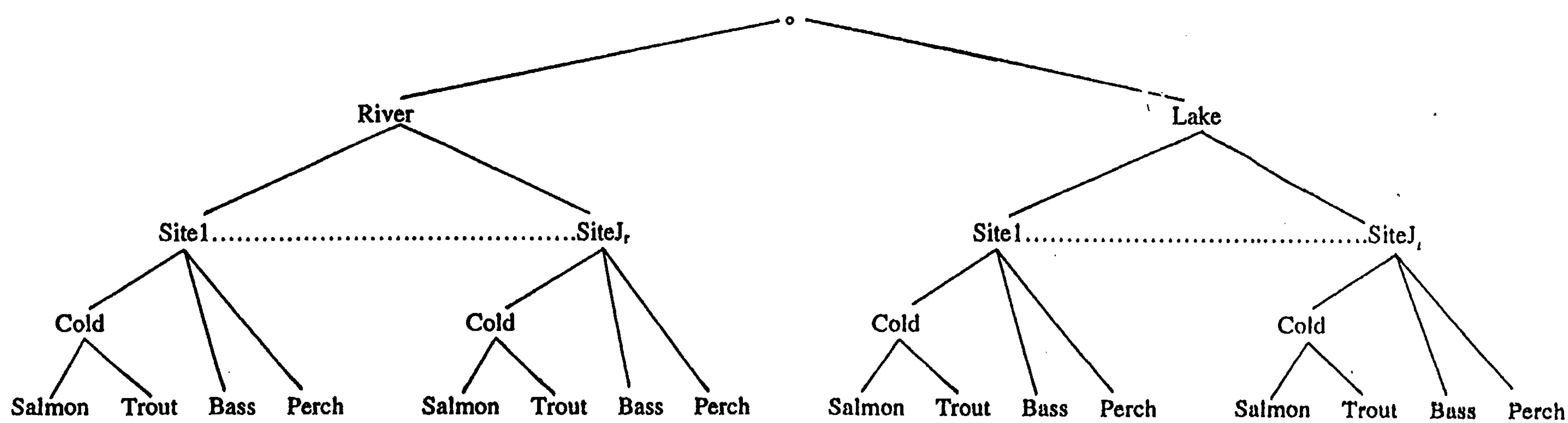


FIGURE 8

4-Level Model (site-species)



Fish Consumption Advisories: Incorporating Angler-Specific Knowledge, Habits, and Catch Rates in a Site Choice Model

Paul M. Jakus, Dimitrios Dadakas, and J. Mark Fly

This study uses a multinomial logit (MNL) site choice model to examine the impacts of sportfishing consumption advisories in eastern Tennessee. The model differentiates by type of angler and by whether or not the angler knew about fish consumption advisories. Further, the estimation method follows Morey and Waldman to endogenously determine catch rates at each fishing site, thus avoiding the biases associated with ad hoc assumptions regarding an angler's catch rate at sites he or she did not actually visit.

Fish consumption advisories are often used to warn recreational anglers that toxic contaminants in fish can result in acute or chronic illness if eaten. Advisories are an

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important management tool because adverse health consequences can be averted while avoiding potentially large clean-up costs. But this management approach assumes that anglers know about advisories and follow recommended practices concerning consumption. In light of a recent study by Holland and Wessells finding that food safety was a key product attribute for fresh seafood, it is not unreasonable for policymakers to assume that safety is an important attribute for sport-anglers catching freshwater fish.[1] Unfortunately there has been little research investigating (a) the assumption that anglers respond to advisories or (b) the costs associated with angler response.

Economists have only recently addressed the issue of fish consumption advisories in the published literature. MacDonald and Boyle found that 63% of anglers in Maine knew about the statewide mercury contamination advisory on lakes and ponds, but fewer than one-quarter of knowledgeable anglers engaged in any behavior designed to protect against contamination (e.g., consume fewer fish or none at all, or fish uncontaminated waters). Among Maine anglers modifying behavior in response to advisories, the seasonal loss in consumer surplus was $151. Jakus et al. estimated a repeated discrete choice travel cost demand model capturing only the site-substitution response of anglers. Seasonal welfare losses associated with a substitution response to advisories were found to be about $47 for anglers in East Tennessee, where fish in six of fourteen major reservoirs were under consumption advisories due to PCB contamination.[2]

The Jakus et al. study represents the only published indirect valuation approach to modeling the impacts of consumption advisories, but the researchers were forced to assume that all anglers were aware of advisories. MacDonald and Boyle, along with a number of other authors (e.g., May and Burger; Diana, Bisogni, and Gall) have cast doubt on this assumption by showing that not all anglers know about advisories,. Further, advisories may have different impacts depending on the angler's goal: those fishing primarily to eat their catch may respond to an advisory differently from a catch and release (C&R) angler who will not eat the catch. It is entirely possible that reduced harvest by consumption anglers may actually increase the site quality for C&R anglers as the stock of fish increases. In this way the welfare changes associated with fish consumption advisories may be positive or negative, depending on angler type.

Data

The University of Tennessee Human Dimensions Lab collected data in the Spring and Fall of 1997 using a random digit dial telephone survey of households from the general Tennessee population. Each survey began with 10,000 randomly selected phone numbers (purchased from Survey Sampling, Inc.). After adjusting for ineligible numbers (disconnects, fax machines, businesses) and numbers at which no contact was made with a household representative, the response rates were 43.5% (Spring 1997) and 47.1% (Fall 1997). Participants were asked about fishing and hunting activities over the six month period immediately prior to the interview.[3] If a respondent indicated he or she had fished in reservoirs, detailed questions were asked about which reservoirs were fished, how often, and the average catch rate (an aggregate of all species) at each reservoir. Respondents were also asked if they fished primarily for C&R or to eat most of their catch and if they knew of fish consumption advisories on Tennessee reservoirs.

A sample of 222 reservoir anglers from a 35 county region of east Tennessee provided complete data for use in the analysis. Anglers in the sample took an average of 14.3 reservoir fishing trips. About 60.8% of the sample said they fished primarily for C&R, while 22.5% fished primarily for consumption of their catch. The remaining 16.7% said they engaged in both C&R and consumption. Just under 65% said they were knew of fish consumption advisories on Tennessee reservoirs.

There are twelve major reservoirs within the 35 county region. Similar biennial surveys (1993 - 1996) conducted by the UT Human Dimensions Lab indicate that anglers also visit a few reservoirs outside the region, but the 1997 sample showed little activity at these reservoirs except for two (Dale Hollow, TN and Hiwassee, NC). The fourteen reservoirs in the choice set accounted for just over 98% of all reservoir fishing trips. The maximum driving time between any origin within the region and any reservoir was less than four hours.[4] Six of the fourteen reservoirs were under some form of fish consumption advisory due to PCB contamination (Boone, Ft. Loudon, Melton Hill, Nickajack, Tellico, and Watts Bar). Advisories ranged from limited consumption of selected species (e.g., less than 1.2 pounds per month of catfish) to an advisory indicating zero consumption of selected species (e.g., no consumption of catfish or striped bass).

Methods

The basic form of the MNL site choice model is reasonably well-known and can be found in any number of publications (e.g., Morey; Bockstael, McConnell, and Strand). It is assumed that on each choice occasion the angler will visit the site yielding the greatest utility. For any two sites *j* and *k*, angler *i* will choose site *j* if the utility at site *j* is greater than the utility at any other site *k*, as in,

$$V_i^j(p_i^j, q_i^j) + \varepsilon^j > V_i^k(p_i^k, q_i^k) + \varepsilon^k \quad \textit{for all } j \neq k$$

where $V(\bullet)$ is the indirect utility function, p_i^j is the travel cost of person *i* to site *j*, q_i^j is the quality experienced by person *i* at site *j*, and all other arguments have been suppressed. Assuming the errors are distributed according to an extreme value distribution, the probability that person *i* visits site *j* can be given by,

$$\pi_i^j = \frac{\exp[V_i^j(p_i^j, q_i^j)}{\sum_k^K \exp[V_i^k(p_i^k, q_i^k)}$$

The log likelihood function is weighted by trips made to each site k by each person i, t_i^k,

$$\ln L = \sum_i^N \sum_k^K t_i^k \ln \pi_i^k$$

and then summed over all people and sites in the sample,

Maximization of (3) yields parameter estimates for the indirect utility function.

This formulation includes a key feature complicating the model: the q_i^j indicate that site quality characteristics may vary with each angler, so that angler specific quality measures are needed for each site. While this does not pose a problem for exogenous site characteristics (e.g., the number of boat ramps), it is a problem for characteristics which may be endogenous to the angler, such as the catch rate at a site, because anglers rarely visit all sites in a choice set. Some measure of "expected" catch is needed.

A common approach is to substitute the mean catch rate for the site, but Morey and Waldman (MW) have recently shown this ad hoc solution results in an errors-in-variables problem because mean catch rates are subject to sampling variability. They demonstrate that the parameters on catch rate and travel cost are biased downward, thus affecting subsequent welfare measures.

This issue has been addressed empirically by Englin, Lambert, and Shaw (ELS) and MW. ELS linked a poisson catch rate model to a poisson aggregate trips model such that both models were estimated simultaneously. MW linked poisson catch rate models for each site to a nested logit model, again where the models are estimated simultaneously. A key difference between the two approaches is that MW estimated the catch rate for each site, while ELS estimated a single catch rate function which varied across sites only as explanatory variables varied across sites.[5]

The MW model uses observed catch rates for each site to measure the probability

$$P(C_i^k) = \frac{\exp(-C^{k^*})(C^{k^*})^{C_i^k}}{C_i^k!}$$

of catch rate per unit effort,

where C_i^k is the observed catch rate for person *i* at site *k* and C^{k*} is the expected catch rate for site *k* as estimated with a poisson process given in (4). It is assumed that errors associated with the trip making process are uncorrelated with the errors from each site, an assumption also made under the ad hoc approach. This assumption is valid if fishing skill and/or practice, as they affect the catch rate, are not site specific. The likelihood function can then be augmented with a poisson catch model for each of the *K* sites. Letting f_i^k equal one if angler *i* visited site *k* and zero otherwise, the log likelihood function making

$$\ln L = \sum_i^N \sum_k^K t_i^k \ln \pi_i^k + f_i^k \ln P(C_i^k)$$

expected catch rates endogenous is given by,

Empirical Results

The catch rate models given by (4) were estimated using only a constant as an explanatory variable, so that the parameter for each reservoir corresponded to the natural log of the catch rate.[6] At each iteration in the estimation process the poisson parameters were converted to catch rates for use in the site choice model. The effect of catch rate (*Catch*) on site choice may vary with angler practice, so *Catch* was interacted with *Boat*, a zero-one dummy variable indicating whether the angler fished mostly from a boat (1) or the bank (0). Empirical models appear in Table 1. White's generalized covariance matrix was used, providing robust standard errors (White).[7]

The first fourteen parameters of each model were the poisson parameters for each reservoir. The estimates were positive and statistically significant in all cases. Also, the *Travel Cost* parameter was negative and statistically significant in all models, as expected. The number of boat ramps at a site (*Ramps*, a measure of site accessibility) was negative in all models (contrary to expectations) but was insignificant.

The first model examined the impact of consumption advisories on site selection, regardless of angler knowledge of advisories or angler type (C&R vs. Consumption). The model assumed that all anglers were aware of advisories. *Catch* was not significant at conventional levels, but the *Catch*Boat* interaction term was statistically significant. The negative sign on *Advisory* indicated that reservoirs with fish consumption advisories were less likely to be visited relative to reservoirs without advisories, all else equal. *Advisory* was not significant with a two-tailed test, but was significant with a one-tailed test of the hypothesis that the coefficient is negative.

Model #1 featured a potentially unpalatable assumption: all anglers were assumed to know about advisories. But an angler who did not know of an advisory would be unlikely to respond to it. In fact, only 65% of the anglers in the sample were aware of advisories. Knowledge of advisories can be cast within the context of the "information" problem found in the contingent valuation literature (e.g., Cameron and Englin). The site quality variable q_i^j capturing the effects of a consumption advisory may be a function of whether the angler was aware of the advisory, so that $q_i^j = q(A^j, K_i)$, where A^j indicates an advisory on reservoir j and K_i indicates knowledge by person i.[8] This was modeled using an interaction of two zero-one dummy variables indicating presence of an advisory (*Advisory*) and angler knowledge (*Know*) of the advisory.

In Model #2 *Advisory**Know* took the value of one if the reservoir had an advisory and the angler had knowledge of advisories. This variable had a value of zero for all anglers who were not aware of advisories. Thus, this specification more closely resembled the information set available to anglers.[9] All coefficients retained the same signs and levels of significance as in Model #1, but the *Advisory**Know* variable was statistically insignificant. This result could have occured for at least two reasons: (a) that advisories do not result in substitution of "clean" reservoirs for "dirty" reservoirs, and so *Advisory* in model #1 captures effects other than those intended or (b) if the *Advisory**Know* variable masks effects that differ across types of anglers.

For example, if an advisory caused consumption anglers to reduce harvest, then as the stock of fish increased a more attractive fishery for C&R anglers may have resulted.

Thus an advisory may have had a negative effect on site selection for consumption anglers and a positive effect for C&R anglers. This hypothesis was investigated in models #3 and #4. In model #3 the dummy variable *Advisory*Know*Consumption* took the value one for consumption anglers who were aware of advisories and zero otherwise. If consumption anglers (22.5% of the sample) engaged in site substitution, a negative sign was expected. In model #3 the sign on this variable is significant and negative, indicating a site substitution response by consumption anglers relative to C&R anglers and anglers who do both types of fishing. In model #4 the variable *Advisory*Know*Catch&Release* took value of one for C&R anglers who were aware of advisories and zero otherwise. Under the preceding hypothesis the expected sign was positive. In fact the sign of this variable was positive but insignificant, while the sign for consumption anglers remained negative and significant.

Welfare gains and losses (under the assumption that all six reservoirs were cleaned up such that advisories could be removed) were estimated for each model. For model #1, which assumed that all anglers are aware of advisories, the per trip welfare gain was $7.29 (confidence intervals reported in Table 1). Model #2 restricted the impact to only those anglers with knowledge of advisories, finding average gain across all anglers to be $1.49.[10] In model #3 average gains for all anglers were $0.38 whereas in model #4 the average per trip losses to C&R anglers outweighed gains to consumption anglers, so that the overall average welfare change was negative $0.25. Mean gains to consumption anglers were $2.33 per trip, while mean losses to C&R anglers were $1.91.

Conclusions

This study has examined the impact of fish consumption advisories, controlling for anglers' knowledge of advisories, the type of angler (consumption vs. catch & release), and endogeneity of catch rates. Both knowledge and angler type have been found to influence the empirical models. Anglers who knew of the advisories and who fished primarily to consume their catch were responsive to the advisories, substituting away from reservoirs with advisories and toward reservoirs without advisories. An opposite effect was expected for C&R anglers: while the coefficient on advisories was indeed positive, it was statistically insignificant. Information about advisories also appeared to play an important role in the travel cost models, indicating a need for research similar to the ongoing research on the role that commodity experience and information play in direct valuation methods.

ENDNOTES

1. The authors used conjoint analysis to examine consumer preferences for three attributes of fresh salmon: farm-raised or wild-caught, price, and inspected by a federal agency or not inspected. The strongest preferences were related to food safety, where inspected products were preferred to those not inspected.
2. Two recent studies using repeated discrete choice travel cost models have examined the general problem of toxics in water. Montgomery and Needelman found that per capita losses due to contamination of New York state lakes and ponds was approximately $63 per year. Hauber and Parsons found that Maine anglers would benefit about $289 per year if all toxic rivers within a four hour drive were cleaned up.
3. The Spring 1997 survey period was September 1, 1996 through February 28, 1997, whereas the Fall 1997 period was March 1, 1997 through August 28, 1997.
4. Parsons and Hauber have shown that recreational sites at this distance for one day trips have a negligible effect on parameter estimates, although other authors provide evidence to the contrary (e.g., Peters et al.).
5. The key complication with the ELS approach is that the errors between the catch model and the travel cost model may be correlated, especially if factors affecting catch rate also affect trip-making behavior.
6. The unit of effort defined by the data is a day of fishing.
7. Models were also evaluated using the inverse of the information matrix. All

parameter estimates (except *Advisory**Aware* in model #3) were statistically significant using this estimate of the variance-covariance matrix. Asymptotically, the White matrix is equivalent to the inverse of the information matrix. If the two differ substantially, as in this study, the White matrix is preferred.

8. While Cameron and Englin have investigated models in which information about the environmental commodity is endogenous, in this study knowledge of advisories is treated as exogenous.

9. This specification does not come without cost. The implied assumption is that the factors giving rise to an advisory (e.g., PCBs) are not perceived in any way by anglers who are unaware of the advisory.

10. Anglers with no knowledge of advisories had no welfare gain or loss under the policy scenario.

Table 1. Multinomial Site Choice Models (222 observations)

Variable	#1	#2	#3	#4
ln Catch (Boone Reservoir)	1.53* (10.93)[a]	1.47* (9.22)	1.49* (8.41)	1.49* (8.55)
ln Catch (Cherokee R.)	1.71* (19.09)	1.73* (18.54)	1.74* (18.03)	1.74* (19.31)
ln Catch (Chickamauga R.)	1.78* (21.09)	1.81* (20.81)	1.81* (21.20)	1.81* (21.36)
ln Catch (Dale Hollow R.)	1.72* (15.52)	1.76* (15.40)	1.75* (15.60)	1.75* (15.87)
ln Catch (Douglas R.)	1.73* (19.48)	1.76* (19.13)	1.76* (19.19)	1.76* (19.35)
ln Catch (Ft. Loudon R.)	1.69* (17.10)	1.65* (17.51)	1.64* (16.58)	1.64* (16.34)
ln Catch (Hiwassee R.)	1.52* (10.79)	1.57* (10.92)	1.59* (11.03)	1.59* (11.52)
ln Catch (Melton Hill R.)	1.51* (10.06)	1.46* (9.13)	1.46* (8.30)	1.46* (8.04)
ln Catch (Nickajack R.)	1.85* (14.73)	1.83* (14.31)	1.80* (13.56)	1.80* (15.04)
ln Catch (Norris R.)	1.80* (21.06)	1.82* (20.41)	1.83* (20.53)	1.83* (20.36)
ln Catch (South Holston R.)	1.45* (9.49)	1.48* (9.42)	1.50* (8.54)	1.51* (8.97)
ln Catch (Tellico R.)	1.78* (17.49)	1.76* (17.66)	1.74* (17.55)	1.73* (19.06)
ln Catch (Watauga R.)	1.42* (8.55)	1.44* (8.34)	1.46* (7.40)	1.47* (7.98)

ln Catch (Watts Bar R.)		1.93* (17.39)	1.90* (18.46)	1.88* (18.00)	1.88* (18.67)
Travel Cost		-0.047* (-8.65)	-0.047* (-8.75)	-0.047* (-8.61)	-0.047* (-8.64)
Ramps	-0.020 (-0.83)	-0.016 (-0.69)	-0.019 (-0.67)	-0.020 (-0.74)	
Catch		1.022 (1.35)	0.948 (1.34)	1.069 (1.12)	1.112 (1.19)
Catch*Boat		0.861* (1.75)	0.825* (1.70)	0.853 (1.60)	0.867 (1.61)
Advisory		-0.722 (-1.42)			
Advisory*Know			-0.228 (-0.59)	0.118 (0.23)	
Advisory*Know*Consumption				-0.856* (-1.69)	-0.700* (-1.74)
Advisory*Know*Catch&Release					0.293 (0.62)
Mean WTP (Clean/Remove all advisories)		$7.29	$1.49	$0.38	-$0.25
95% Confidence Interval[b]		-$2.60 - $22.75	-$3.10 - $6.69	$0.22 - $0.59	-$0.47 - -$0.11

* statistically significant at $\alpha=0.10$ (two-tailed test).

[a] Number in parentheses is the ratio of a coefficient to its asymptotic standard error.

[b] Calculated using the method of Krinsky and Robb.

References

Bockstael, N.E., K.E. McConnell, and I. Strand. 1991. "Recreation." In *Measuring the Demand for Environmental Quality*, eds. J.B. Braden and C.D. Kolstad, chapter 8, New York: North-Holland.

Cameron, T.A. and J. Englin. 1997. "Respondent Experience and Contingent Valuation of Environmental Goods." *J. Environmental Economics and Management*, 33(3):296-313.

Diana, S.C., C.A. Bisogni, and K.L. Gall. 1993. "Understanding Anglers' Practices Related to Health Advisories for Sport-Caught Fish." *J. Nutrition Education*, 25(6):320-28.

Englin, J., D. Lambert, and W.D. Shaw. 1997. "A Structural Equations Approach to Modeling Consumptive Recreation Demand." *J. Environmental Economics and Management* 33(1):33-43.

Holland, D. and C.R. Wessells. 1998. "Predicting Consumer Preferences for Fresh Salmon: The Influence of Safety Inspection and Production Method Attributes." *Agricultural and Resource Economics Review*, 27(1):1-14.

Jakus, P.M., M. Downing, M.S. Bevelhimer, and J. Mark Fly. 1997. "Do Sportfish consumption Advisories Affect Reservoir Anglers' Site Choice?" *Agricultural and Resource Economics Review*, 26(2):196-204.

Krinsky, I. and A.L. Robb. 1986. "On Approximating the Statistical Properties of Elasticities." *Review of Economics and Statistics* 68(4):715-19.

MacDonald, H.F. and K.J. Boyle. 1997. "Effect of a Statewide Sport Fish Consumption Advisory on Open-Water Fishing in Maine." *North American J. Fisheries Management*, 17:687-95.

May, H. and J. Burger. 1996. "Fishing in a Polluted Estuary: Fishing Behavior, Fish Consumption, and Potential Risk." *Risk Analysis*, 16(4):459-71.

Montgomery, M. and M. Needleman. 1997. "The Welfare Effects of Toxic Contamination in Freshwater Fish." *Land Economics*, 77(3):211-23.

Morey, E.R. 1998. "Two RUMs unCLOAKED: Nested Logit Models of Site Choice and Nested-Logit Models of Participation and Site Choice." In *Valuing the Environment Using Recreation Demand Models*, eds. C.L. Kling and J. Herriges, chapter 4. Northampton, MA: Edward Elgar Publishing, Ltd.

Morey, E.R. and D.M. Waldman. 1997. "Measurement Error in Recreation Demand Models: The Joint Estimation of Participation, Site Choice, and Site Characteristics." Department of Economics, University of Colorado (October)

Parsons, G. R. and A.B. Hauber. 1998. "Spatial Boundaries and Choice Set Definition in a Random Utility Model of Recreation Demand." *Land Economics*, 74(1):32-48.

Peters, T., W.L. Adamowicz, and P.C. Boxall. 1995. "Influence of Choice Set Considerations in Modeling the Benefits of Improved Water Quality." *Water Resources Research* 31:1781-87.

White, H. 1980. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica* 48:817-38.

An Analysis of Wildlife Recreation Using the FHWAR

by

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An Analysis of Wildlife Recreation Using the FHWAR

The Conservation Reserve Program (CRP) may effect wildlife populations, with resulting impacts on public participation in non-consumptive wildlife based outdoor recreation. To study this possible relationship, we use data from U.S. Fish and Wildlife Service's Fishing, Hunting, and Wildlife Associated Recreation (FHWAR, US Fish and Wildlife Service) survey. This is a survey of approximately 50,000 indivduals nationwide (pared down from a 250,000 person screener survey). Approximately 26,000 individuals were asked about their non-consumptive wildlife associated recreation (i.e.; wildlife viewing) activities. In particular, we focus on the number of trips taken in the general vicinity of the individual's home; which roughly translates to all trips taken within 100 miles of the individual's residence.

The size and extent of the FHWAR database are the primary features motivating its use. However, counterbalancing these positive features is the paucity of information about the recreational sites visited by respondents. For our analysis, site specific information is very important, since we are most interested in how landscape characteristics (i.e.; the extent of CRP) influence recreational behavior.

To deal with this lack of information, a representative trips model is used. This entails modeling the aggregate trip taking behavior, of an individual, as a function of the characteristics of sites available to her. Lacking good information on which site was visited, we let the data guide the model.

Briefly, the analysis involves the following steps:

I) Using GIS tools, create several "landscape characteristics" variables defined at a number of "potential visitation zones" around each respondent. These are then aggregated, using a flexible weighting scheme, into "distance-zone-aggregated" landscape charcteristic variables (**Z**)

II) For each individual, extract visitation (**Q**), and socoioeconomic (**X**), data from the 1991 FHWAR.[1].

III) Using the "distance to most frequently visited site" as a dependent variable, estimate a representative trip price (**P**).

IV) Regress total number of trips against the **X** and **Z**, and on **X, P** and **Z**

V) Using coefficients from step IV, estimate predicted number of trips (and consumer surplus) under several scenarios.

The following sections provides greater detail on each of these steps.

I) Imputing Landscape Characteristics.

An Average Shift Histogram (technique is used to impute landscape characteristics. These characteristics, which are primarily based on National Reources Inventory (NRI) data, include:.

i) %CRP. The percent of the land (in the sub-county region) that is in CRP.

ii) %CROP.

iii) %FOREST.

iv) %GRASSAND (rangeland and pasture).

v) RUC: Rural Urban Continuum code (0 being most urban, 9 being most rural).

vi) DIVERSITY: Landscape diversity, with higher values of DIVERSITY indicating a more variegated landscape. This diversity is based on the extent and interspersion of water, forest, grassland, and cropland.

For each FHWAR respondent, a set of 19 "potential visitation zones" are defined. These zones are drawn from five distance bands surrounding the respondent's zip code. In particular:

Zones	Description	
1 :	The "own zip code" zone.	
2 to 4 :	18 miles to origin	the center of each zone is seperated by 120 degrees.
5 to 7 :	37 miles to origin	the center of each zone is seperated by 120 degrees.
8 to 12 :	62 miles to origin	the center of each zone is seperated by 72 degrees.
13 to 19 :	87 miles to origin	center of each zone is seperated by 51 degrees.

Note that zones 5 to 19 are larger then zones 1 to 4. Also, for each of these 19 "potential visitation zones" the several landscape characteristic variables are obtained.

To simplify the model (and avoid problems with missing data[2]), the landscape characteristics of the 19 potential visitation zones are aggregated into five "distance zone" measures (**Z**). The simplest aggregation would consist of averaging the value of (a a given landscape characteristic) across all "visitation zones" within a " distance zone". However, to deal with the possiblity that landscape heterogeniety may be important, we use a "constant elasticity of substitution" (*CES*) functional form to compute an aggregate measure.

Specifically, this measure is defined as:

$$\text{i) } Z_k = \left(\sum_{j=1}^{J_k} \left(LC_{jk} \right)^{1/\alpha} \right)^{\alpha}$$

where:

α is a parameter to be estimated,
J_k is the number of zones in the k'th *distance* band (i.e.; band four has 5 zones),
LC_{jk} is the value of the characteristic in the j zone of band k

and

Z_k is the **band** measure of the land characteristic.

Note that when
$\alpha = 0$: Maximum matters
$\alpha = 1$: Sum (or average) matters
$\alpha >> 1$: Variations in characteristics don't matter

Given the K=5 distance-zones, and six characteristics (listed in section I), the above process yields 30 seperate aggregated landscape characteristic (**Z**) variables.[3]

II) Individual Data

Data on non-consumptive wildlife associated recreation was obtained from the 1991 FHWAR. For each surveyed individual, number of visits to "non-distant" sites was extracted. Operationally, this involved several steps:

1) Using information on past participation, and on current plans, observations on individuals who were not likely to be "potential wildlife viewers" were dropped .[4]

2) Trips to one's own state, and to all states for which the "most visited location" within a 100 miles of the resident's home, were summed to obtain total "non-distant" trips (for all "potential wildlife viewers").[5]

3) Several socioeconomic variables were extracted for each individual, including Male, Caucasian, Rural Residence, High School Graduate, College Graduate and Household Income. All but Household Income are dummy (0/1) variables. Household income is an approximation based on the center of broad ranges (in the $0 to >$75,000 interval). In cases where income was not available, a zip-code average was used (the average was generated from the 250,000 respondents interviewed in the screener portion of the FHWAR).

4) Individual weights were also obtained for each observation. These are demographics weights, computed by the FHWAR survey designers, that are used when creating population level predictions.

IV) The Model

Using the X (socioeconomic), P (imputed price), and **Z** (aggregated landscape characteristics) a "representative trip" demand curve is estimated.

To clarify, lacking good information on where people went, our model focuses on the total number of wildlife associated trips within a few hours drive of an individual's residence. Hence, the use of the landscape characteristic (**Z**) variables to estimate total trips is interpreted as a highly reduced form of the site-selection problem solved by an individual recreator. That is, the model combines trips to all sites into a "total number of trips", and uses aggregated landscape characteristics (with the aggregation occurring over all sites) to explain the total number of trips taken. Thus, the determinants of a chosen trip (to an unknown-to-the-analyst site) is "represented" by these aggregated characteristics.

To control for the prevalence of zero trips, a double hurdle Poisson estimator is used to model this representative trip model.

iii)
$$Prob(q; q>0 \mid \lambda,\gamma) = \left(\frac{e^{-\lambda}\lambda^{q}}{q!}\right)\left(1-e^{-\gamma}\right)$$

$$Prob(q=0 \mid \lambda,\gamma) = e^{-\lambda} + \left((1-e^{-\lambda})e^{-\gamma}\right)$$

where:

λ is the quantity parameter: $\lambda = \exp(R\beta)$
γ is the the participation parameter: $\gamma = \exp(S\tau)$
β and τ are parameters to estimate.
R are factors that influence the number of trips; including P (price information), X (socioeconomic factors), and Z (aggregated landscape characteristics)
S are factors that influence participation (typically, a subset of X is used)

In it's simplest form, equation iii does not incorporate explicit price information. Instead, variations in the quality and price are controlled through the use of the distance zone specific variables. For several reasons, it would be desirable to include explicit price information. First, if explicit price information can be obtained (say, a "representative" price), then welfare estimates using consumer surplus may be readily computed. Second, the inclusion of such price information should improve the performance of the model.

The problem is, as with landscape characteristics, the paucity of knowledge about which sites were visited implies a lack of explicit price information; a problem that is exacerbated when individuals took zero trips.

As a substitute, a predicted "representative" price can be used.

The predicted "representative" price is based on the FHWAR distance to most frequently visited site variable. This distance variable is converted into a dollar cost, using average cost per mile information, and a simple time cost (based on a fraction of household income). This price is then used as the dependent variable in a sample selection model. The use of this predicted price offers two advantages: as a control for potential simultaneity between "price" and "quantity of trips", and to impute a price for the (many) individuals who consumed zero trips.

To predict this price, a sample selection model is used:

iv)

$$1)\ Prob[Participant;\gamma] = \Phi(\chi,\gamma)$$

$$2)\ E[Price;\beta_\chi,\beta_v \mid Q>0] = \chi\beta_\chi + \frac{\phi(\chi,\bar{\gamma})}{\Phi(\chi,\bar{\gamma})}\beta_v$$

where :

$\chi = X \text{ and } Z \text{ variables}$

$\bar{\gamma} = \text{Predicted value of } \gamma \text{ from step 1}$

$\phi = \text{Normal pdf}$

$\Phi = \text{Normal cdf}$

The first equation is a Probit on whether the individual took any trips, with Z and X used as regressors. The coefficients from the Probit(γ) are used to compute a Mills ratio. This Mills ratio, along with the Z and X variables, are regressed against the log of observed price (using observations with non-zero trips) in a standard semi-log OLS. Lastly, the predicted values of γ,β_χ,and β_γ are used with equation iv.2 to compute a predicted price (E[P;·) for all observations.

IVa) Estimating the Model

The model to be estimated is:

Z = Z(LC,α)
P = P(X_1,Z;γ)
q = F(X_1,X_2,Z,P,W;β_1,β_2,β_p,β_z)

where:

q= Number of trips.
X_1= Individual specific variables that influence probability of participation; typically socioeconomic variables.
X_2= Individual specific variables that influence quantity of trips; typically socioeconomic variables. Note that X_1 and X_2 may contain the same variables.
Z = Aggregated landscape characteristic variables, for z different variables and k=1..K bands. These will be a function of the 19 LC variables (each LC variable is specific to a segment of a distance band around an indivdidual) and α (the "CES" aggregation parameter).
P = The imputed price of a trip. Based on a sample selection model with the observed "distance to favorite site" as the dependent variable, and X_1 and Z as the independent variables.[6]
Z() = The "distance zone" variables aggregation model (equation i)

P() = The "sample selection" model (equation iv)
F() = The hurdle Poisson model (equation iii)
W = Population weight correction factor.[7]

and

$\beta_1,\beta_2,\beta_p,\beta_z,\alpha,\gamma$ = Parameters to be estimated.

Although simultaneous estimation of the above would be optimal, operational difficulties dictate a multi-stage model; to wit:

1) Using a grid search, select a candidate value of α.
   a) For each candiate value of α, the Z_k variables are generated.
   b) Given Z, P (prices) are then imputed.
   c) Given Z and P, estimate F(·).
   d) Record the log-likelihood from c.
2) Reiterate step 1 for different values of α.

3) Given a set of coefficient vectors (one vector for each value of α) choose the one with the best log-likelihood. The β coefficients associated with this best log-likelhood are the estimated parameters of the model.

Since it might be expected that recreational behavior may vary across the country, this model was applied seperately to the five sub-national regions: the West, Northern Plains, Southern Plains, North, and the South.[8]

2) Constructing alternative scenarios

Given the estmated coefficients, we then predict total recreational trips under different allocations of the CRP. We focus on 3 scenarios:

1) 1991_CRP: The CRP as of measured by the NRI; totalling about 34 million acres.[9] Note that the estimation of the model is based on this scenario.
2 15_EBI: 34 million acres, but using 15th EBI to select acres
3) NO_CRP: No CRP

Scenario 1 is based on an allocation of NRI points to the CRP as a function of expected costs and landscape characteristics; with the total acreage of CRP maintained at the level reported by the 1992 NRI (about 34 million acres). However, the distribution of CRP lands across the nation does change.
The NO_CRP scenario reallocates 1991 CRP to alternative land uses, such as cropland and rangeland (this reallocation uses information contained in the 1982 NRI).

Table I lists the %CRP (and % CROP) "percieved" by the FHWAR sample. Since the FHWAR sample is not uniformly distributed geographically, these percentages will differ from the actual landscape distribution in the regions.

Percieved %CRP and %CROP (average over 100 mile region), 3 Scenarios and 5 Regions

	1991_CRP (34 Million NRI Acres)		15_EBI		No CRP	
	%CRP	%CROP	%CRP	%CROP	%CRP	%CROP
West	1.2	9.9	1.1	9.9	0	11.2
N. Plains	4.5	52.3	4.3	52.4	0	56.9
S. Plains	0.9	19.2	1.4	18.6	0	20.0
North	1.1	28.9	1.6	27.8	0	29.9
South	0.9	14.6	1.3	13.6	0	15.4

** The percents are calculated against all land within 100 miles of an individual's residence, and averaged over all individuals in the sample. To the extent that individuals in a region are not uniformly distributed, these "percieved" values will differ from the actual distribution of land use.

Given an alternative scenario, the Z variables for each observation will be different. The impacts of these changes is examined by recomputing the predicted number of trips, using the recomputed aggregated landscape characteristics (Z), X, P, and the estimated coefficients from the model. Ideally, the coefficient on the imputed price would then be used to estimate a total consumer surplus for the sample, and a U.S population consumer surplus (using the FHWAR weights).

3) Some Results

Across the nation, approximately 2/3 of the sample were classified as being potentially interested in "non-consumptive wildlife associated" recreation. About one half of these individuals (1/3 of the sample) actually took at least one trip. The remaining third were classified as uninterested, and were not included in the estimatation. The following table gives further details:

Regional summary of participation in non-consumptive wildlife associated recreation

	# Obs	# Retained	# Participants	Average #trips	Average distance to visited site	Average Reported Trip Value[a]
West	5561	3391	1624	9.33 (16.7)	22 (33)	30
N. Plains	2075	1679	659	11 (23)	13 (18)	25
S. Plains	992	785	270	8.9 (19)	25 (44)	31
North	9827	7878	3122	13 (23)	14 (30)	32
South	6451	4699	1547	10 (19)	15 (19)	31

[a]the average reported trip value is based on a contingent valuation question asked of all non-consumptive trip takers.

The canonical estimator for this model, as described in section IVa, is based on a double hurdle Poisson model and an imputed price. Given the large number of variables, the following table lists some of the more important variables. The "sum" of the β_z coefficients for each landscape characteristic is also displayed; which can be interpreted as the effect given a uniform change in landscape characteristics.

Note that the coefficients are best interpreted as the percent change given a unit change in the variable. The probability variables range from 0 to 100; the RUC ranges from 0 to 9, and the diversity variable ranges from 1 to 4.

Some coefficients from the double hurdle model with imputed price (t-stats in parenthesis).

	West	N.Plains	S. Plains	North	South
Aggregation Parameter					
α	0.47	2.33	1.5	4.7	0.47
Some probability stage coefficients					
INCOME	-2.13e-7 (-.19)	-1.58e6 (-0.82)	4.14e-6 (1.5)	-1.02 (-1.3)	1.22e-6 (1.0)
OWN_CRP	-0.019 (-12.1)	-0.052 (-2.3)	0.13 (3.4)	0.042 (4.1)	-0.047 (-1.66)
Some quantity stage coefficients					
Income	-9.34e-6 (-10.7)	5.13e-6 (8.4)	6.1e-6 (4.4)	-9.5 (-15.5)	-5.0e-6 (-6.5)
CRP01	0.352 (12.2)	0.023 (0.61)	-1.17 (-5.8)	0.011 (1.17)	0.89 (17.4)
CRP02	-0.234 (-9.9)	-0.002 (-0.44)	0.30 (5.7)	0.00053 (7.7)	-0.65 (-15.4)
CRP03	0.0077 (0.86)	0.0010 (0.40)	-0.13 (-4.7)	-0.00066 (-9.0)	0.15 (8.3)
CRP04	0.061 (12.8)	0.0013 (2.1)	-0.004 (-0.42)	-4.31 (-0.5)	0.016 (1.9)
CRP05	-0.028 (-8.3)	-8.06e-5 (-0.4)	0.037 (-5.60)	1.6e-5 (10.0)	-0.003 (-0.056)
Price	0.029 (4.8)	-0.0093 (-0.90)	-0.002 (-0.34)	0.133 (11.9)	0.027 (3.14)
Summation of landscape characteristic coefficients					
∑ %CRP	0.16	0.02	-0.96	0.01	0.40
∑ %CROP	0.57	.10	0.02	-0.01	-0.001
∑ %Forest	0.15	.27	-0.04	0.003	0.008
∑ %Grass	0.005	.03	-0.01	-0.03	0.02
∑ %Diversity	-0.31	-2.58	.37	-0.02	1.6
∑ RUC	0.11	0.05831	0.29	0.06	-0.003
Log Likelihood	17671	10733	3239	59802	20316

These coefficients are somewhat difficult to interpret, as they show no strong pattern. %CRP seems to be more often positive than negative, with the exception of the S. Plains.

Given the goals of this study, it is of especial interest to examine how trip taking behavior changes under the various scenarios. The following table lists the actual (Q_p), and the *population-weighted* prediction (Q*_p), total number of trips under the 3 scenarios. Where available, the estimated Consumer Surplus is also displayed.

Non-consumptive Wildlife Related Recreation
Hurdle Poisson Model using Population Weight and Imputed Price
Total # of trips (in 1,000s) and *CS (in $1,000)*

	Observed *Average $CS*	Original *$CS*	S1 (15th ebi) *$CS*	No CRP
West	43,047	46,177 *n.a.*	47,469	47,343
N. Plains	4,815 *108*	4,907 528,956	4,907 428,917	3,837 413,360
S. Plains	9,958 *499*	10,169 *??*	11,868	8,158
North	105,938	113,023 *n.a.*	117,502	104,001
South	36,049	40,662 *n.a.*	42,356	40,504
Total	199,807	214,938	222,715	203,846

Notes:

i) *n.a.* means the $CS was not computable due to a positive sign on the price coefficient.

ii) ?? means that a small in magnitude price coefficient yielded a $CS that was implausibly large (i.e.; $500 or more per trip).

iii) The observed number of trips and the predicted number of trips (using the original data) are not necessarily equal.

As a measure of model quality, the correlation between the weighted observed and weighted predicted number of trips (based on the original scenario) can be used in lieu of an R-square statistic.

	Imputed Price Model	
	Correlation: Individual trips	Correlation: weighted trips
West	0.18	0.40*
N. Plains	0.31	0.15
S. Plains	0.43	0.41
North	0.24	0.18
South	0.19	0.17

*When a large outlier was not removed from the West, the correlation was 0.81.

Ideally, the imputed price model would yield immediately useful consumer surplus values. Unfortunately, the price coefficient values are often positive (or negative but very small in magnitude); which yields impossible (or implausible) consumer surplus values. It would appear that the distribution of quality sites obscures the price relationship.[10] However, since the "imputed price" does allow extra information (the "distance to last site" data) to be incorporated, we will retain the results with the understanding that the "price" is to be interpreted loosely.[11]

As an alternative to a directly computed consumer surplus, we can use a benefit's transfer value. In particular, the "average per day" value of wildlife watching can be used as a proxy for per-trip value. Although several sources for such a value exist, the "self-reported" value from the FHWAR is most appropriate for this exercise. The regional averages of these values are used to report the "consumer surplus" of wildlife viewing trips under the three scenarios.

	Total CS ($1,000)	**West**	**N. Plains**	**S. Plains**	**North`**	**South**
Original	6,700,482	1,385,310	122,675	315,239	3,616,736	1,260,522
15_EBI	6,987,753	1,424,070	122,676	367,907	3,760,064	1,313,036
NO_CRP	6,352,769	1,420,290	95,925	252,898	3,328,032	1,255,624

Examining the TOTAL CS column, the national impact, on benefits due to wildlife related recreation, of the 15th EBI is approximately $288 million dollars. This is in addition to a baseline benefit (attributable to the existence of the CRP) of approximately $347 million dollars. In other words, the redistribution of acreage, without any increase in the size of the program, yielded an approximately 82% increase in the impact of the CRP on wildlife related recreation. Other points include:

The bulk of the increase in benefits due to the 15th EBI occur in the North and the South; a result of the larger changes in CRP combined with large populations.

* The Southern Plains accounts for a relatively large increase (proportional to it's baseline value)
* The N. Plains is essentially stable, which is not suprising (given that the N. Plains looses CRP acreage under the 15th EBI.
* The West has a relatively small change. Somewhat disturbing is the positive (albeit small) impact associated with the NO_CRP scenario.

4) Summary

This paper presents a methodology that allows one to estimate the effects of landscape attributes on the demand for wildlife related recreation. In particular, the model introduces a representative trips model. This model uses aggregate measures of landuses as a proxy for information on the quality of recreational sites available to individuals. To increase the flexibilty of the estimator, a flexible weighting function is used to construct these aggregate measures.

Using GIS techniques and NRI data, wildlife trip data from the 1991 FHWAR was used to estimate the representative trip model. Several alternative scenarios, based on different distributions of the CRP, were then simulated, and changes in expected trip demand were computed. These simulations suggest that use of the "15th EBI" would lead to a fairly large (approximately $300 million) increase in consumer surplus.

The major strength of this modeling effort is that it uses a large dataset (the FHWAR) in a microeconometric model of demand. However, it does suffer from the need to use an aggregation function (albeit one that is sensitive to the data). Furthermore, the use of landuse measures as proxies for site quality is driven out of necessity, and is far from ideal. Future work on the valuation of CRP should address these weaknesses.

ENDNOTES.

References

Scott, David W. and Gerald Whittaker, "Multivariate Applications of the ASH Regression," *Communications in Statistics*, 25(11), 1996: 2521-30.

U.S. Department of the Interior, Fish and Wildlife Service and U.S. Department of Commerce, Bureau of the Census. 1991 National Survey of Fishing, Hunting, and Wildlife Associated Recreation. U.S. Government Printing Office, Washington DC, 1993.

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1. This requires knowledge of a key piece of information: the individual's residence. Since public-use releases of FHWAR do not contain this information (due to privacy concerns), analysis of this data necessitated use of raw data at U.S. Census facilities.

2. For example, due to potential visitation zones that lie in the ocean, or some other large water body.

3. The five distance zones could be further aggregated into an overall measure by using an endogenous distance decay Although this yields a more parsimoniously specified model, it also complicates estimation.

4. The following table contains the percent of observations in 4 categories.

	0 Trips	>0 Trips
Excluded Observation	18%	7%
Included Observation	45%	30%

Notes:

Ideally, the "excluded-observation/>0 trips" category should contain 0% (since individuals who took a trip should not be a priori excluded).

Conversely, the "included-observation/0 Trips" category may contain a large percentage, since it is possible for potential participants to choose 0 trips in any given season.

Approximately 85% of all trips were accounted for by individuals retained in the sample.

5. This focus on trips to "non-distant" sites is necessitated by modeling concerns; such as the large number of "sites" one would have to include in order to account for far away trips. However, note that trips to these "non-distant" sites account for over 90% of non-consumptive wildlife associated trips.

6. The price term is computed as the sum of an out of pocket cost and a time cost:

P= [0.3 * DIST] + [WAGE * 0.33 (DIST/50)]

Where,

DIST = Distance to site (in miles)
0.3 = Approximate per mile cost of using a car
WAGE = Imputed wage rate = Household income divided by 2040.

DIST/50 = Time required to travel DISTANCE
0.33 = Fraction of travel time that is "onerous." The assumption is that recreational travel is not as unpleasant as work, hence should not be valued at the wage rage.

Note that several assumptions are made, including:

i) The WAGE rate assumes that the trip taker is the soul wage earner in the household; and freely chooses to work 2040 hours.

ii) Out-of-pocket costs (0.3 * DIST) assumes a group size of one (no cost sharing, and no variation in fuel economy, depreciation rates, etc.

iii) An average speed of 50 miles per hour.

7. When using the FHWAR weights to scale up to the population, the desired equivalence between "observed" and "predicted" (using the baseline data) number of trips need not hold. There are several ways of addressing this inconsistency; including ex-post calibration, weighted estimation, or inclusion of the weight as a correction factor. Though all of these are problemmatic, the use of the weight as a correction factor involves the fewest ad-hoc assumptions.

8. The five regions consist of:

1) West: CA, WA, OR, MT ID WY NV UT CO AZ NM
2) N. Plains SD ND NE KS
3) S. Plains OK TX
4) NorthEast: MN WI MI IA MO IL IN OH ME VT NH CT RI MA NY PA MD NJ DE DC
5) SouthEast: AR LA MS AL GA SC FL KY TN SV VA NC

9. This model assumes that the landscape at the time of the 1991 FHWAR is well represented by the 1992 NRI. There are two means by which this representation may be inaccurate. First, the sample of points covered by the 1992 NRI accounts for about 34 milion acres of CRP, out of a total 1991 CRP acreage (from signups 1 to 11) of about 35.4 million acres. That is, the NRI undercounts the CRP by about 1.4 million acres. Second, the 1.5 million acres signed up in 1991 (signups 10 and 11) probably did not effect recreational behavior (given that cover crops, etc. had not been established). Given that these two sources of error are roughly equivalent, and are fairly small, the 1992 NRI data is used as is.

10. If the distribution of site quality varies over the population (with some individuals living close to better sites, while others must travel long distances to attain better sites), then the imputed price should be correlated with number of trips. That is, better quality sites nearby should yield more trips to close in sites; hence a negative sign on the imputed price coefficient. On the other hand, if the shape of the distribution of site quality is similar across the population (say, increasing with respect to distance), but with some individuals having better all around choices (say, the slope of the distance/quality relationship varies across individuals), then high prices may be associated with high number of trips. That is, individuals who can pay a high price for the "best sites" may take more trips then individuals who choose a closer in "mediocre" site.

11. A number of other specifications were attempted, including models without imputed price terms, and models that used the simple Poisson model. The results from these models were qualitatively similar to the double hurdle, imputed "price" model.

Random Digit Dialing versus Retired Sample From the Current Population Survey:

The case of the National Survey of Fishing, Hunting, and Wildlife-Associated Recreation.

By

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1. Introduction.

The 1996 National Survey of Fishing, Hunting, and Wildlife-Associated Recreation (Survey) was the ninth in a series of national surveys that collected data on wildlife-related recreation. The 1996 Survey collected data on hunting, fishing, and wildlife watching (observing, photographing, and feeding) for the 1996 calendar year. This is the most comprehensive survey of it kind that collects wildlife-related recreation data on a state-by-state basis.

The comprehensive and complex nature of the Survey has lead to major increases in the cost of the Survey. The cost of the Survey reached 13 million in 1991. The high cost results from sampling and interviewing techniques including the use of the Bureau of Census' retired CPS sample, in-person follow-up interviews to respondents not reached by phone, and the trimester recall that requires contacting the respondents three times throughout the year.

Begining in 1991 changes were made to reduce the cost of the survey. The first big change was to use primarily telephone interviewing instead of in-person. Since then, several other changes have been made to reduce the cost of the Survey. These include:

Reducing the sample size
Reusing the base sample from 1991 while adding for new construction.
Combining the screening and first detailed interview (sampling on the fly).
Interviewing respondents in two waves instead of three (counting on correct answers from household respondents during screening).
Multiple phone attempts before using personal interviewing.
Dropping personal interviews for the middle wave and picking up non-interview in the final wave.
Removing layering in management of survey (Team approach)

These changes, and others, resulted in a substantial savings. The 1996 Survey cost $7.5 million, a 50% cut in real terms from the 1991 cost. This amounts to $150,000 per state. This is the relevant figure since the Survey is truly 50 independent surveys that sum to a national survey. That is, 50 independent samples are selected to obtain reliable estimates for each state.

We are continuing to look for ways to reduce the cost of the Survey.

The Fish and Wildlife Service is currently looking at the possibility of using Random Digit Dialing as a sampling method instead of the retired Current Population Survey Census based sample.

The results presented here are from research conducted to test for inherent differences in the two sampling methods. This is especially important because of the importance put on trends in fishing, hunting, and wildlife watching.

This is a summary of results. A final report can be obtained from Sylvia Cabrera, Division of Federal Aid, USFWS, 4401 N. Fairfax Drive, RM 140, Arlington, VA 22203.

Sampling definitions

Current Population Survey

The CPS is a regularly conducted survey that collects information that is used to update the decennial data on population, income, employment, etc.

The CPS is a stratified sample that represents the nation and is representative down to the state level.

Sample persons are used in the CPS for a maximum of 2 years, after which they are "retired".

Once retired, they are sometimes called back and used for special surveys like the fishing and hunting survey.

Random Digit Dialing

The Genesys program was used to select sample for the 1996 Survey RDD sample.

This program randomly selects phone number banks (100 banks) from all banks with at least one working number in order to get complete geographic coverage. (E.G. you would have (111) 111-11xx)

Phone numbers are randomly selected from the "100 banks" by randomly selecting 2 digits to complete the number.

Interviewing procedures for the RDD and CPS samples were nearly identical. For the RDD sample, phone number were matched with addresses using Telematch. Letters were sent to respondents in both sample alerting them that they would be contacted about their fishing and hunting activities.

The big difference between the samples is that interviewers attempted personal contacts for CPS sample persons who were not reached by phone.

Table 1 shows a brief comparison of the sample on some demographic characteristics. The samples are very similar — with income being the only exception.

Table 1. Demographic Comparisons Between CPS and RDD Sample

	RDD	CPS
Sex		
Male	47.8%	48.0%
Female	52.2%	52%
Race		
White	82.3%	83.1%
Black	9.2%	9.3%
Other	8.5%	7.6%
Ethnicity		
Non-Hispanic	93.0%	92.8%
Hispanic	7.0%	7.2%
Household Income		
Less than $25,000	29.3%	34.6%

Differences of less than 1% are not significant at the 0.05 level

Table 2 shows estimates for fishing participation.

Table 2. Comparison of Participation Estimates from RDD vs. CPS Samples For Fishing: 1996			
Anglers and Days	RDD	CPS	% Change
Anglers (thousands)			
Total	40,248	35,246	14.2%
Freshwater	33,483	29,734	12.6%
Freshwater, except Great	32,562	28,921	12.6%
Great Lakes	2,414	2,039	*18.4%
Saltwater	12,089	9,438	28.1%
Days of fishing (thousands)			
Total	743,111	625,893	*18.7%
Freshwater	632,710	515,115	22.8%
Freshwater, except Great	581,918	485,474	*19.9%
Great Lakes	30,041	20,095	Not Significant
Saltwater	114,115	103,034	Not Significant
Average days per angler			
Total	18.5	17.8	Not Significant
Freshwater	18.9	17.3	Not Significant
Freshwater, except Great	17.9	16.8	Not Significant
Great Lakes	12.4	9.9	Not Significant
Saltwater	9.4	10.9	Not Significant

Percents in the "Percent Change" column marked with an asterisk are significant at the 0.10 level. All other Percents are significant at the 0.05 level.

Table 3 shows participation estimates for Hunting.

Table 3. Comparison of Participation Estimates from RDD vs. CPS Samples For Hunting: 1996			
Hunters, days, and	RDD	CPS	% Change
Hunters (thousands)			
Total	16,642	13,975	19.1%
Big Game	13,478	11,288	19.4%
Small Game	8,249	6,945	18.8%
Migratory Bird	3,830	3,073	24.6%
Other animals	1,589	1,521	Not Significant
Days of hunting (thousands)			
Total	323,253	256,676	*25.9%
Big Game	197,845	153,784	28.7%
Small Game	89,322	75,117	Not Significant
Migratory Bird	32,870	26,501	Not Significant
Other animals	31,342	24,522	Not Significant
Average days per hunter			
Total	19.4	18.4	Not Significant
Big Game	14.7	13.6	Not Significant
Small Game	10.8	10.8	Not Significant
Migratory Bird	8.6	8.6	Not Significant
Other animals	19.7	16.1	Not Significant

Percents in the "Percent Change" column marked with an asterisk are significant at the 0.10 level. All other percents are significant at the 0.05 level.

Findings on participation generally show a significant difference between estimates from the two samples. And the estimates are greater for the RDD sample in each case.

The next two tables show comparisons of expenditure estimates for the two different samples.

Table 4. Comparison of Expenditure Estimates from RDD vs. CPS Samples For Fishing: 1996			
Expenditures	RDD	CPS	% Change
Fishing expenditures			
Total	$59,497,602	$37,797,061	57.4%
Freshwater	$35,908,048	$24,482,439	46.7%
Freshwater, except Great	$34,341,215	$22,445,123	53.0%
Great Lakes	$1,345,562	$1,404,885	Not Significant
Saltwater	$15,205,027	$8,081,276	88.2%
Average expenditure per angler			
Total	$1,478	$1,072	37.9%
Freshwater	$1,072	$823	30.2%
Freshwater, except Great	$1,055	$776	35.9%
Great Lakes	$557	$689	Not Significant
Saltwater	$1,258	$856	46.9%

Percents in the "Percent Change" column marked with an asterisk are significant at the 0.10 level. All other percents are significant at the 0.05 level.

Hunting expenditure estimates from the two samples are shown in Table 5.

Table 5. Comparison of Expenditure Estimates from RDD vs. CPS Samples for Hunting: 1996			
Expenditures	RDD	CPS	% Change
Expenditures (thousands)			
Total	$28,261,502	$20,613,412	*37.1%
Big Game	$12,933,281	$9,712,735	Not Significant
Small Game	$3,810,052	$2,481,385	*53.5%
Migratory Bird	$1,980,755	$1,296,322	Not Significant
Other animals	$607,964	$432,593	Not Significant
Average expenditures per			
Total	$1,698	$1,475	Not Significant
Big Game	$960	$860	Not Significant
Small Game	$462	$357	Not Significant
Migratory Bird	$517	$422	Not Significant
Other animals	$383	$284	Not Significant

Percents in the "Percent Change" column marked with an asterisk are significant at the 0.10 level. All other percents are significant at the 0.05 level.

Summary

The CPS sample base which had been tried and tested and is assumed to be comprehensive in it's representation is consistently showing lower estimate for participation and expenditures for fishing and hunting.

Differences in participation range from 12% to 28%.

Differences in expenditures range from 37% to 88%.

So, what does all this mean for the National Survey of Fishing, Hunting, and Wildlife-Associated Recreation?

> We will be planning for the 2001 survey and this information will be presented to the data users and they can express their thoughts.
>
> Since trend information is so important, it is unlikely that an RDD survey will be used unless it is supplemented in some way to compensate the sample. One possibility is to use some CPS sample to augment the RDD survey and adjust for any bias associated with the RDD sample. Another possibility is to supplement a CPS sample with RDD to reduce sampling costs but maintain sample integrity.

What does this mean for other surveys that are done using RDD?

> We can't really generalize to all surveys but RDD based surveys of hunters and anglers may be overestimating the actual level of participation and expenditures.

What are other options?

> We are still doing comparisons on RDD vs. Telephone only CPS. In this comparison we will only compare the RDD estimates to the responses from the CPS sample that were reached by phone. If the differences disappear, it will give further support that the CPS sample base is superior. However if the differences are still there, is possible that the bias associated with RDD is not just RDD based but telephone based. That is, the CPS sample is better because households not reached by phone are contacted in person. If this is the case, another solution would be to use a RDD sample base and follow up with personal interviews. (Addresses can be obtained by matching phone number with addresses.)

INTEGRATED CONTINGENT VALUATION - TRAVEL COST MODELS FOR EXPENDITURE ALLOCATION DECISIONS

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INTEGRATED CONTINGENT VALUATION - TRAVEL COST MODELS FOR EXPENDITURE ALLOCATION DECISIONS

I. INTRODUCTION

In many nonmarket valuation studies, data for applying more than one economic valuation or economic impact analysis technique are collected in the same survey instrument. An example is found in a study of the economic values and impacts of alternative reservoir water level reported by Cordell and Bergstrom. In this study the survey instrument collected data necessary for estimating economic values based on the contingent valuation method (CVM) and the travel cost method (TCM). Expenditure data were also collected for estimating economic impacts via input-output analysis or some other appropriate economic impact analysis technique.

It has been a common practice in previous studies that have collected data for applying both stated and revealed preference nonmarket valuation techniques for researchers to estimate separate models based on these techniques and then make value comparisons (e.g., Smith, Desvousges, and Fisher). In recent years, researchers have explored simultaneous estimation of revealed and stated preference values using integrated models (Cameron). The primary purpose of this article is to propose and test a theoretical model and data analysis techniques for integrating CVM, TCM, and expenditure data to estimate economic values for nonmarket environmental commodities. The theoretical model and data analysis techniques will be applied to estimation of the recreational value of alternative reservoir levels.

We begin in the next section by developing a model of discrete-continuous choice in a recreation management program that is consistent with utility theory. We demonstrate that a two-equation model may be derived from a theoretically valid utility difference model commonly used for describing dichotomous choices for contingent valuation (CV) scenarios. We then apply this model to the reservoir water level example to demonstrate the feasibility of explicitly linking the two decisions through the utility-consistent derivation.

II. WELFARE MEASURES FROM DICHOTOMOUS CHOICE MODELS

Consider the general case of estimating use value associated with an individual's use of some recreational area such as a forest, beach, or reservoir. Hanemann (1984) derives Hicksian compensated measures from a utility difference model estimated from dichotomous choice, contingent valuation data. The observed discrete choice response of each individual is assumed to reflect a utility maximization process.

We assume that utility can be decomposed into an observable component, V_i, where the subscript denotes the two states, and an unobservable component, ε_i. The unobservable factors that influence an individual's indirect utility function are independent and identically distributed random variables with zero mean.

In contingent valuation studies, welfare changes are typically and appropriately defined in terms of an annual lump-sum income decrement interpreted as annual willingness-to-pay (WTP). The annual lump-sum income decrement or annual WTP corresponds to a Hicksian compensating or equivalent welfare measure. In recreation studies, an annual lump-sum income decrement associated with access to a recreation area may be defined in terms of payments in special taxes to keep the area open or perhaps an annual recreational fee for access.

The observable portion of the indirect utility function for each individual depends on both the annual lump-sum income change to an individual of participating in recreation activities at the recreational area, *A*, and the variable costs of participating in the recreation experience, *P*. The annual cost of access to a recreational area is in addition to household income, *Y* for the respondent regardless of the level of participation decision. Characteristics of the recreation experience and the respondents socioeconomic background, *Z*, also affect the participation decision and thus indirect utility.

An individual will participate in recreation at a recreational area if $V_1(Y-A, P, Z)$ is greater than $V_0(Y, Z)$, where the indirect utility of participation is denoted by V_1 and that of nonparticipation by V_0. The probability of participation in the pass program is defined as

$$Pr(\,Participation\,) = Pr\left[\,V_1(\,Y+P\,,A\,,Z\,)+\varepsilon_1 > V_0(\,Y\,,Z\,)+\varepsilon_0\,\right]$$
$$= Pr\left[\,\varepsilon^* < V_1(\,Y+P\,,A\,,Z\,) - V_0(\,Y\,,Z\,)\,\right]$$

where all terms are as previously defined. The unobserved factors that influence responses to the CV scenario are represented by ε^*, the difference in the error terms of the indirect utility functions, defined as ε_0 - ε_1. The error term ε^* is treated as a single random variable assumed to follow a logistic distribution.

The Recreation Trip Decision

Individuals who agree to incur the annual costs of access to a recreational area, say in the form of an annual pass, must then decide on the number of trips to make to the area during the next 12 months. The trip decision is a continuous choice that flows directly from the utility function underlying the recreation participation decision. A model to analyze these two decisions must account for the theoretical and empirical linkages between the discrete and continuous choices.

Using duality theory, we derive the decision on the number of trips from the same indirect utility function that generates the discrete response from the dichotomous choice, contingent valuation question. This results in a theoretically consistent two-equation model of decision behavior. We apply an econometric method for estimating discrete-continuous choice models that corrects for selectivity bias.

McFadden and Leonard and McConnell demonstrate that the utility difference model can be used to derive the ordinary demand for the good identified in the CV scenario using Roy's Identity. McConnell used the offer payment to represent the income term in Roy's Identity, since for the linear difference model, marginal utility of income is constant and independent of state.

$$X^* = \frac{\partial \Delta V}{\partial P} / \frac{\partial \Delta V}{\partial A}$$

Using Roy's Identity the quantity demanded can be written as where *P* represents the price of the good and *A* is the bid amount in the CV scenario and the specific form of the demand function depends on the utility difference equation specified.

<u>A Utility-Theoretic Specification for Discrete-Continuous Choices</u>

$$\Delta V = \alpha_0 + \beta_1 \ln(Y - A) + \beta_2 \ln P + \gamma \ln Z + \varepsilon^* .$$

We specify a loglinear approximation to the underlying indirect utility difference model for the decision to incur annual costs for access to a recreational area:

Variables that influence the decision to pay for continued access to a recreational area but do not affect the number of recreation trips, are denoted by *Z*.

The price of each trip is estimated by each recreationist and is hypothesized to be influenced by the type of trip planned by the recreationist along with socioeconomic characteristics. The systematic variation in the coefficient for the price or

$$\beta_2 = \delta_0 + \delta_1 \ln P + \delta_2 \ln W \ .$$

expenses on each trip is modelled as

Here *W* is a vector of variables representing factors related to the recreationist's determination of trip costs, including human and physical capital (such as boats, camping equipment) invested in previous outdoor recreation trips and information about recreation area conditions. Behaviorally, this specification implies that respondents use trip costs and information contained in *W* to choose the number of trips that is optimal based on the recreation experiences for that household.

$$\Delta V = \alpha_0 + \beta_1 \frac{A}{Y} + \delta_0 \ln P + \delta_1 \ln P \ln P + \delta_2 \ln P \ln W + \gamma \ln Z + \varepsilon^* \ .$$

We substitute equation [4] into equation [3] and recall that the loglinear utility model implies a specific form of the bid and income variable in the utility difference model:

If $\Delta V > 0$, the respondent chooses to incur the annual costs of continued access to a recreational area.

The utility difference model yields the logit specification when the probability of a yes response is specified as the cumulative distribution function of a standard logistic variate. Mean WTP from the utility difference model is calculated following Hanemann (1989).

The number of trips taken decision depends on the utility

$$X = \frac{(\delta_0 + 2*\delta_1 \ln P + \delta_2 \ln W)/P}{\beta_1 / Y}$$

difference equation and is derived using Roy's Identity as:
Note that the second stage decision be formulated as the total

$$\frac{X*P}{Y} = \omega_0 + \omega_1 \ln P + \omega_2 \ln W + \mu$$

expenditures on trips as a proportion of household income: where X represents the number of trips and P denotes the expenses incurred on each trip and the parameters are defined as $\omega_0 = \frac{\delta_0}{\beta_1}$, $\omega_1 = \frac{2\delta_1}{\beta_1}$ and $\omega_2 = \frac{\delta_2}{\beta_1}$. The error term μ, which incorporates the effect of unobserved factors on the trip decision, is assumed to be normally distributed with zero mean in the population of all recreationists.

This specification also controls for heteroscedasticity arising when variation in the number of planned recreation trips depends on the household's income. Even after accounting for differences in household characteristics and types of recreation experiences, the number of trips may vary with the total income for each household.

Equations [5] and [7] compose the utility-theoretic discrete-continuous choice model. The linkage between the two choice is due to the fact that both decisions arise from a single underlying utility function. Information on the number of planned trips is observed only for those individuals who agree to incur the annual lump-sum costs of continued access to a recreation area. A form of selectivity bias arises and can be corrected for in the estimation.

Correcting Selectivity Bias

Selectivity bias is due to the correlation between ε^* from the discrete decision and μ from the continuous decision, implying that $E(\varepsilon^*\mu) \neq 0$. Since the subsample of participants used to estimate the continuous decision is selected from the sample of all respondents, unobservable variables that affect the utility a recreationist derives from participation outdoor recreation may influence the decision on the proportion of expenditures allotted to recreation trips. This correlation means that the conditional mean of μ is not zero, given the decision is made to participate and pay annual lump-sum costs of access. Due to this sample selection bias, parameter estimates based on ordinary least squares (OLS) are biased and inconsistent.

The estimation technique accounts for sample selection bias that is present when uncorrected OLS is used for the planned trip decision. Under the assumptions that ε^* follows a logistic distribution and μ is normally distributed in the population of all recreationists with zero mean and variance $\sigma 2$, and representing the correlation between ε^* and μ in the population as $\rho_{\mu\varepsilon}$, Dubin and McFadden showed that

$$E(\mu|\Delta V>0)=\kappa T,$$

$$\kappa=-\frac{\sqrt{6\sigma_\mu^2}}{\pi}\bullet\rho_{\mu\varepsilon}$$

$$T=\left[\frac{(1-Pr)\ln(1-Pr)}{Pr}+\ln Pr\right] \quad .$$

where

The conditionality term *T* increases in *Pr*, the estimated probability that the recreationist agrees to participate and pay annual lump-sum costs. The greater the probability of participation, the larger the conditionality term for a respondent. The coefficient κ is positive if the unobserved factors in the participation decision are negatively correlated with the total trip expenditures, that is if $\rho_{\mu\varepsilon} < 0$. The error term μ can be decomposed into the conditional mean and deviations and substituted into equation [7] to yield:

$$\frac{X*P}{Y}=\omega_0+\omega_1\ln P+\omega_2\ln W+\kappa T+\mu^* \quad .$$

Including κT permits unbiased estimates of the parameters of the model as well as unbiased prediction of the percentage of recreation expenditures spent on trips, conditional on willingness to incur annual lump-sum costs of access (say in the form of an annual recreational pass).

III. EMPIRICAL APPLICATION

The general model described in the previous section was applied to the task of estimating the recreational values of alternative reservoir water levels using an integrated CVM-TCM model. The study area as described in more detail by Cordell and Bergstrom was four Tennessee Valley Authority reservoirs located in western North Carolina and north Georgia. Management of these reservoirs for flood control and hydropower involves relatively large fluctuations in water levels.

Water level drawdown for flood control and hydropower directly affects reservoir use for outdoor recreation. Low water levels in summer and early fall reduce the suitability of reservoirs for recreational activities such as motor boating, sailing, water skiing, swimming, and fishing. A study was conducted to collect data for estimating the recreational value and impacts of alternative reservoir water levels. Economic valuation results based on the CV data collected are reported and described in Cordell and Bergstrom.

In the CV questionnaire described by Cordell and Bergstrom, recreationists were asked to suppose that everyone using a reservoir would be required to purchase an annual recreation pass for that reservoir. After describing the payment vehicle, bids for the current management situation were elicited by a dichotomous choice or referendum approach. Recreationists were asked if they would continue to use the reservoir if the annual recreation pass were required at a given posted-price. Posted prices ranged from $1.00 to $300.00. The range of posted prices was based on previous literature and a pre-test of the survey questionnaire.

Respondents were asked to provide their name and address in order to receive follow-up mail questionnaires on expenditure and equipment usage on recreation visits. The information gathered in the questionnaire consisted of expenditure variables.

Respondents provided trip-related expenditures for specific items within the general categories of food, lodging, transportation, activities and miscellaneous. These specific items represented the major purchases of most recreation visitors and corresponded to IMPLAN economic sectors. For each item, the survey asked the amount spent at home in preparation for the trip or after returning, while traveling to or from the visited site, and at or in the immediate vicinity of the site during the trip.

IV. ECONOMETRIC SPECIFICATION

A summary description of the variables used in the estimation along with sample means and standard deviations is presented in Table 1. The study area focuses on four reservoirs in western North Carolina and North Georgia: Lake Chatuge, Lake Fontana, Lake Hiwassee, and Lake Santeetlah, which were indicated by a set of lake dummy variables (LAKE1, LAKE2, LAKE3, and LAKE4). The explanatory variables for each household include the annual income of the respondent (INC), the number of years of education (EDUC), the sex of the respondent (SEX), and the number of years the respondent has been using lakes in the western North Carolina area for outdoor recreation (YEARS).

Data from the equipment and expenditure survey were used to form two variables reflecting allocation of expenses between on-site expenses and off-site expenses incurred preparing for the trip or when travelling to and from the lake site. Expenses at the lake site are denoted by EXPAT. Expenses for the trip incurred at home in preparation for the trip or after returning or while traveling to or from the visited site are denoted by EXPPREP. In the empirical application of the discrete-continuous choice model, the second stage model consists of two decisions modelled jointly: on-site expenditures and trip preparation or off-site expenses.

Let the vector W_1 represent the on-site expenses for the recreation trip (EXPAT) and a set of interaction terms relating to equipment usage on the trip. Data on equipment usage included the total number days a camping vehicle or tent was used (CAMPTOT), the total number days a boat was used (BOATTOT), and total fees such as admission fees, parking fees, or entrance fees for the recreation trip (TOTFEES). We use W_2 to represent the trip preparation expenses for the recreation trip with the same set of interaction terms relating to equipment usage on the trip.

Empirical Results for the Discrete-Continuous Choice Model

Maximum likelihood coefficient estimates of the empirical utility difference model are presented in Table 2 which shows the complete set of variables included in the utility difference model. Table 2 is arranged in three columns to highlight household specific variables which influence only the discrete choice (column 1) and the interaction terms relating to trip allocation expenses which also influence the continuous, second-stage choice (columns 2 and 3).

The negative coefficient on the bid amount relative to income in the utility difference model indicates that higher bid amounts (as a proportion of total income) decrease the probability of a yes response. Each of the reservoir identification variables has a negative sign and are statistically significant and these variables are suppressed in the table. The statistically significant negative signs suggest that the economic value of higher water levels is lower at the three lakes relative to Lake Chatuge (the "base" reservoir).

The continuous choice allocation model of trip expenditures given participation in the pass program is estimated as a seemingly unrelated regression model. Coefficient estimates for the two equations are presented in Table 3.

These expenditure models incorporated the conditionality terms, CTERM1 AND CTERM2, given as κ in equation 9. The CTERM was significant and negative in the preparation expenditures model. The unobserved factors that affect the decision to pay fees and participate in the recreation activities, ε^* in equation 5, are positively correlated with the allocation of preparation expenditures for recreation activities through the term $\rho_{\varepsilon\mu}$. This is reflected in the definition of κ given in equation 9.

The last recreationists agreeing to accept the CV scenario have the least to gain from participating, exhibiting lower positive changes in utility, ΔV from equation 5, and reflecting smaller values of ε^*. In turn, the values of unobserved factors μ that influence the expenditure decisions are smaller, reflecting the positive correlation between ε^* and μ. As the marginal recreationist decides to participate, the average percentage of preparation expenditures provided by this recreationist decreases.

Failure to apply the self-selectivity corrections produces misleading assessments of the key variables which influence expenditures. Hensher and Milthorpe emphasized the importance of selectivity-based models, noting that even if the selectivity variable is not statistically significant, "its inclusion is necessary to detect and account for the magnitude of selection bias on *individual* parameters."

V. CONCLUSIONS

A behavioral model for combining stated and revealed preference nonmarket valuation techniques is proposed. The same utility function should theoretically govern responses to both types of valuation methods. Yet, researchers may overlook methods to construct a utility-consistent framework which exploits restrictions derived from rational economic choices in empirical analysis of survey data. The empirical application of the discrete-continuous choice model consists of two continuous decisions modelled jointly: on-site expenditures and trip preparation or off-site expenses. The two-equation construct permits specification and testing of different factors that may influence the two expenditure decisions.

Our study demonstrates the benefits of the discrete-continuous choice method: that it generates a consistent behavioral model that uses all information available from a survey, produces theoretically and econometrically linked discrete and continuous decision models, permits correction of selectivity bias, and produces a welfare measure in which policy makers may place more confidence than those evolved from *ad hoc* models.

REFERENCES

Cameron, T.A. "Combining Contingent Valuation and Travel Cost Data for the Valuation of Nonmarket Goods." *Land Economics* 63(1992):302-17.

Cordell, H.K. and J.C. Bergstrom. "Comparison of Recreation Use Values Among Alternative Reservoir Water Level Management Scenarios." *Water Resources Research*. 29 (1993):247-258.

Dubin, J., and D. McFadden. "An Econometric Model of Residential Appliance Holdings and Consumption." *Econometrica* 52 (1984):345-62.

Hanemann, W. M. "Welfare Evaluations in Contingent Valuation Experiments with Discrete Responses." *American Journal of Agricultural Economics* 66 (1984):332-341.

Hanemann, W. M. "Welfare Evaluations in Contingent Valuation Experiments with Discrete Response Data: Comment." *American Journal of Agricultural Economics* 71 (1989):1057-1061.

Hensher, D.A., and F.W. Milthorpe. "Selectivity Correction in Discrete-Continuous Choice Analysis." *Regional Science and Urban Economics* 17(1987):123-150.

McConnell, K. E. "Models for Referendum Data: The Structure of Discrete Choice Models for Contingent Valuation." *Journal of Environmental Economics and Management* 18 (1990):19-35.

McFadden, D. and G.K. Leonard. "Issues in the Contingent Valuation of Environmental Goods: Methodologies for Data Collection and Analysis." in J.A. Hausman (ed.) *Contingent Valuation: A Critical Assessment*. Amsterdam: North-Holland, 1993.

Smith, V.K., W.H. Desvousges, and A. Fisher. "A Comparison of Direct and Indirect Methods for Estimating Environmental Benefits." *American Journal of Agricultural Economics* 68(1986):280-291.

Table 1. Variables Used in Model Estimation

Variable Name	Variable Mnemonic	Mean[a]
Offer price for an annual pass to use a study reservoir (in U.S. dollars)	BID	89.65 (97.90)
Household income for the recreationist (in thousands of U.S. dollars)	INC	33.504 (17.813)
On-site recreation expenditures for the most recent trip to study reservoirs	EXPAT	126.17 (273.51)
Trip preparation expenses for the most recent trip to study reservoirs	EXPPREP	42.15 (79.61)
Total number of days using camping equipment	CAMPTOT	19.33 (27.202)
Total number of days using boating equipment	BOATTOT	19.65 (36.12)
Number of years recreationist has visited the study reservoirs	YEARS	16.923 (12.00)
Number of years of education for recreationist	EDUC	14.54 (3.00)
Sex of respondent (0, female; 1, male)	SEX	0.34 (0.48)
Indicator variables for reservoirs		
Lake Chatuge	LAKE1	0.35
Lake Hiwassee	LAKE2	0.13
Lake Santeetlah	LAKE3	0.42
Lake Fontana	LAKE4	0.10
Access fees paid for the most recent trip to study reservoirs	FEES	15.72 (31.04)
Planned trips with pass access system	PLTRIPS	11.70 (19.04)
Sample Size	N	520

Table 2. Model for Contingent Valuation Participation Decision

	Variable	Variable	Variable
INTERCEPT	5.112* (2.311)		
BID/INC	-0.973* (-7.328)		
YEARS	0.114 (1.364)		
SEX	1.241* (4.120)		
EDUC	-1.777* (-2.483)		
EXPAT		0.124 (1.510)	
EXPAT*EXPAT		-0.007 (-0.525)	
EXPAT*CAMPTOT		-0.004 (-1.287)	
EXPAT*BOATTOT		0.005 (1.623)	
EXPAT*FEES		0.001 (0.432)	
EXPPREP			-0.143 (-1.245)
EXPPREP*EXPPREP			-0.004 (-0.227)
EXPPREP*CAMPTOT			0.002 (0.443)
EXPPREP*BOATTOT			0.0006 (0.240)
EXPPREP*FEES			-0.005 (-1.418)

[a] Asymptotic t-values in parentheses with significance at 0.10 level.

Table 3. Model for Trip Expenditures Given Participation in the Program

	On-Site Expenses	Preparation Expenses
INTERCEPT	-392.74* (-5.292)	-12.402 (-1.478)
EXPAT	71.887* (4.354)	
CAMPTOT	6.805* (8.089)	
BOATTOT	0.690* (1.643)	
FEES	-0.326 (-0.331)	
CTERM1	-13.354 (-0.875)	
EXPPREP		1.555* (2.656)
CAMPTOT		0.256* (2.324)
BOATTOT		0.021 (0.381)
FEES		0.177 (1.370)
CTERM2		-3.900* (-1.922)

[a] Asymptotic t-values in parentheses with significance at 0.10 level.

ROBUST ESTIMATORS OF WILLINGNESS-TO-PAY IN THE DICHOTOMOUS CHOICE CONTINGENT VALUATION FRAMEWORK*

by

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Proposed Running Head: *Robust Estimators of WTP*

The standard procedures for extracting willingness-to-pay (WTP) from dichotomous choice CV questionnaires rely heavily upon parametric assumptions regarding the distribution and form of WTP in the sampled population. However, theory provides little guidance regarding the appropriate parametric specifications to use and the resulting WTP estimates can be quite sensitive to the selections made. In this paper, we compare and contrast the performance of several parametric and semi-nonparametric estimators that have been proposed in the literature, examining the sensitivity of the resulting WTP estimates to the underlying distribution of preferences and the estimation procedure employed.

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I. INTRODUCTION

Welfare analysis in the environmental arena is often complicated by absence of observable market transactions (i.e., revealed preferences) from which to infer the value placed in an environmental good or service. To fill this void, many researchers have turned to the stated preference methods of Contingent Valuation (CV). Dichotomous choice CV, in particular, has come to dominate much of this literature. Within this framework, survey respondents are presented with a hypothetical change in environmental quality and, in the case of a quality improvement, a proposed cost of acquiring the change. The individual's willingness to incur the proposed costs reveals information about the value placed in the environmental improvements. Unfortunately, the standard procedures for extracting the implied willingness-to-pay (WTP) of an individual, as well as the distribution of WTP in a target population, rely heavily upon parametric assumptions regarding the nature of consumer preferences. For example, Cameron's [2] bid function approach begins by segmenting the individual's WTP into two components: (1) a nonstochastic bid function that is assumed to depend upon observed characteristics of the individual and the environmental attributes being valued and (2) a stochastic component or residual used to capture variations in preferences. Typically, researchers then make parametric assumptions regarding both the functional form of the WTP and the distribution of the error term, estimating the model via maximum likelihood techniques. Theory, however, provides us with little guidance regarding the appropriate parametric specifications to use and the resulting WTP estimates can be quite sensitive to the selections made.[1]

The possible bias of parametric estimators has received considerable attention in the general discrete choice literature (e.g., Manski [22], Cosslett [4], Stoker [28], and Matzkin [24]), with studies appearing directly in the CV literature only more recently (e.g., Kriström [21], Chen and Randall [3], and Creel and Loomis [7]). Yet, while a variety of nonparametric and semi-parametric estimators have been proposed, only limited information exists on the gains (or losses) of these estimators relative to the standard parametric procedures, or of the factors that are likely to influence these gains.[2] The purpose of this paper is to partially fill this gap. We contrast the performance of several parametric and nonparametric estimators that have been proposed in the literature using a Monte Carlo framework, examining the sensitivity of the resulting WTP estimates to the underlying distribution of preferences and the estimation procedure employed. In process, we provide an adaptation of the Generalized Maximum Entropy (GME) estimator introduced by Golan, Judge and Perloff [12] to the contingent valuation problem.

The remainder of the paper is divided into six sections. Section II provides a brief overview of the dichotomous choice contingent valuation method and sets up much of the paper's notation. We then describe in Section III the four estimators to be contrasted in our Monte Carlo analysis. These include the parametric probit and log-probit models used extensively in the CV literature, Chen and Randall's [3] semi-nonparametric (SNP) estimator, and an adaptation of the GME estimator of Golan, Judge, and Perloff [12]. The structure of the Monte Carlo exercise is detailed in Section IV, with the results presented in Section V. An application of all four estimators is then presented in Section VI using the same data on water quality valuation employed by Chen and Randall [3]. Section VII provides the conclusions from our analysis.

II. DICHOTMOUS CHOICE CONTINGENT VALUATION

The contingent valuation method relies upon survey questionnaires to elicit information about an individual's evaluation of a nonmarket good or service. While a variety of survey formats have been proposed, the referendum or dichotomous choice format currently dominates the literature. In this setting, survey respondents are presented with hypothetical changes to both an environmental amenity and their level of income. The individual's willingness to accept the income change reveals information about the compensating variation that they associate with the proposed environmental change. This information can in turn be used to conduct welfare analysis.

In order to fix ideas, consider a proposed environmental improvement. Let[3]

$$WTP_i = W(X_i, \varepsilon_i; \beta) \tag{1}$$

denote the i^{th} individual's underlying willingness-to-pay for the environmental improvement, where X_i is a vector of socio-demographic characteristics and β is a vector of unknown coefficients. The disturbance term ε_i is assumed to capture variations in preferences within the population including unobserved individual characteristics. Let B_i denote the corresponding income reduction, or bid, posed in the CV question. One of the advantages of the dichotomous choice format touted in the literature is that it parallels the type of decisions typically made by consumers in the marketplace; i.e., accepting or rejecting a good or service at a fixed price (B_i). The key disadvantage of the format is that the survey response reveals only limited information about the consumer's underlying WTP, bounding above or below the proposed bid. Thus, rather than observing the consumer's WTP, the analyst observes only the latent variable no_i, where

$$no_i = \begin{cases} 1 & W(X_i, \varepsilon_i; \beta) < B_i \\ 0 & W(X_i, \varepsilon_i; \beta) \geq B_i. \end{cases} \tag{2}$$

Discrete choice econometric methods are then brought to bear on the problem in order to characterize the distribution of WTP in the population, rather than the WTP of a given individual. In particular, it is common practice to assume that the ε_i enters the bid function W in an additive fashion, so that

$$W(X_i, \varepsilon_i; \beta) = w(X_i; \beta) + \varepsilon_i, \tag{3}$$

where $w(X_i; \beta)$ denotes the nonstochastic portion of WTP. The analyst then postulates a specific form for the cumulative distribution of ε_i, $\Lambda(\varepsilon_i)$, so that:

$$\begin{aligned} \Pr(no_i = 1) &= \Pr[W(X_i, \varepsilon_i; \beta) < B_i] \\ &= \Pr[\varepsilon_i < B_i - w(X_i; \beta_i)] \\ &= \Lambda[B_i - w(X_i; \beta)]. \end{aligned} \tag{4}$$

The resulting log-likelihood function is given by

$$\mathsf{L} = \sum_{i=1}^{n} no_i \ln\{\Lambda[B_i - w(X_i;\beta)]\} + \sum_{i=1}^{n} (1-no_i)\ln\{1-\Lambda[B_i - w(X_i;\beta)]\}. \tag{5}$$

Maximum likelihood techniques can then applied to estimate the parameters of the model. The problem with the standard parametric approach is that it is not clear what functional forms should be used in specifying either $w(X_i;\beta)$ or $\Lambda(\varepsilon_i)$.

III. ALTERNATIVE ESTIMATORS

A variety of functional forms and estimators have been proposed in the literature for estimating the distribution of WTP from dichotomous choice CV surveys. In this section, we review two parametric and two semi-nonparametric approaches.

Parametric Estimators

Among the most common parametric model employed in the CV literature is the linear probit model. This specification assumes that[4]

$$W(X_i, \varepsilon_{Ni}; \beta_N) = \beta_N' X_i + \varepsilon_{Ni} \tag{6}$$

where $\varepsilon_{Ni} \sim i.i.d.N(0,\sigma_N^2)$ and $w(X_i;\beta) \equiv \beta_N' X_i$. Thus, the probability of a "no" response is:

$$\begin{aligned} \Pr(no_i = 1) &= \Lambda[B_i - \beta_N' X_i] \\ &= \Phi\left[\frac{B_i - \beta_N' X_i}{\sigma_N}\right] \\ &= \Phi[\delta_N' Z_i], \end{aligned} \tag{7}$$

where $\Phi(\cdot)$ denotes the standard normal cdf; $\delta_N = (\delta_{N0}, \delta_{N1}, \ldots, \delta_{Nk})' \equiv (\sigma_N^{-1}, -\sigma_N^{-1}\beta_N')'$; and $Z_i = (B_i, X_i')$ The corresponding log-likelihood is given by:

$$\mathsf{L}_N = \sum_{i=1}^{n} no_i \ln[\Phi(\delta_N' Z_i)] + \sum_{i=1}^{n} (1-no_i)\ln[1-\Phi(\delta_N' Z_i)]. \tag{8}$$

An important attribute of the linear probit model in the CV setting is that, unlike most probit applications, the dispersion of WTP in the population (captured by σ_N) can be separately identified (Cameron [2]). This is accomplished by varying the bids (i.e., the B_i's) across observations. In particular, if $\hat{\delta}_{Nk}$ denotes the k^{th} element of the maximum likelihood estimate of δ_N, then $\hat{\sigma}_N = \hat{\delta}_{N0}^{-1}$. The original parameter vector can likewise be recovered using $\hat{\beta}_{Nk} = \hat{\delta}_{Nk} / \hat{\delta}_{N0}$. Finally, we note that in the probit framework both the conditional mean WTP ($\mu_X \equiv E(WTP|X)$) and the conditional median WTP ($m_X \equiv Median(WTP|X)$) are given by

$$\mu_X = m_X = \beta_N' X_i. \tag{9}$$

The conditional dispersion of WTP in the population is given by

$$d_X \equiv StdDev(WTP|X) = \sigma_N \tag{10}$$

Another commonly employed parametric estimator is the linear log-probit model. Here, it is assumed that the bid function takes the form

$$W(X_i, \varepsilon_{Li}; \beta_L) = \exp(\beta_L' X_i + \varepsilon_{Li}) \tag{11}$$

where $\varepsilon_{Li} \sim i.i.d. N(0, \sigma_L^2)$, or equivalently

$$\ln[W(X_i, \varepsilon_{Li}; \beta_L)] = \beta_L' X_i + \varepsilon_{Li}. \tag{12}$$

The corresponding likelihood is

$$\mathrm{L}_L = \sum_{i=1}^{n} no_i \ln[\Phi(\delta_L' Z_i)] + \sum_{i=1}^{n} (1 - no_i) \ln[1 - \Phi(\delta_L' Z_i)], \tag{13}$$

where $Z_i = [\ln(B_i), X_i']'$. Again, $\hat{\sigma}_L = \hat{\delta}_{L0}^{-1}$ and $\hat{\beta}_{Lk} = \hat{\delta}_{Lk} / \hat{\delta}_{L0}$. In the case of the lognormal specification, the conditional mean WTP is given by:

$$\mu_X = E[\exp(\beta_L' X_i + \varepsilon_{Li})] = \exp\left(\beta_L' X_i + \frac{\sigma_L^2}{2}\right). \tag{14}$$

whereas the conditional median WTP corresponds to $m_X = \exp(\beta_L' X_i)$. Finally, the conditional dispersion of WTP in the population is given by:

$$d_X = \exp(\beta_L' X) \sqrt{\exp(2\sigma_L^2) - \exp(\sigma_L^2)} \tag{15}$$

A Semi-Nonparametric Estimator

A number of authors have recently proposed relaxing the restrictions of the standard parametric models, relying instead on flexible approximations to the unknown distribution preferences. In particular, Chen and Randall [3] have proposed a semi-nonparametric (SNP) estimator for WTP.[5] The authors begin by assuming that the bid function has the structure:

$$W(X_i, \varepsilon_{Si}; \beta_S) = \exp[w(X_i; \beta_S)]\varepsilon_{Si} \equiv G(X_i; \beta_S)\varepsilon_{Si} \tag{16}$$

where $G(X_i; \beta_S)$ is an unknown function characterizing the nonstochastic portion of willingness-to-pay and ε_{Si} is an unknown disturbance term with an unknown distribution. Chen and Randall [3] use the exponential form for $G(X_i; \beta_S)$, together with the restriction that ε_{Si} has support only for nonnegative values, to ensure that willingness-to-pay is nonnegative, i.e. $WTP_i \geq 0$. This structure for the bid function then implies that:

$$\begin{aligned} \Pr[no_i = 1] &= \Pr[G(X_i; \beta_S)\varepsilon_{Si} < B_i] \\ &= \Pr\left[\varepsilon_i < \frac{B_i}{G(X_i; \beta_S)}\right] \\ &= \Lambda\left[\frac{B_i}{G(X_i; \beta_S)}\right] \\ &= \Lambda[u_i] \end{aligned} \tag{17}$$

where

$$u_i \equiv \frac{B_i}{G(X_i; \beta_S)}. \tag{18}$$

In order to reduce the reliance on a specific model parameterization, the authors use flexible approximations to the two unknown functions of the model: $G(\cdot)$ and $\Lambda(\cdot)$.

Gallant's [8] Fourier Flexible Form (FFF) is used to model the nonstochastic portion of the individual's bid function. That is, $w(X; \beta_S)$ is approximated by:[6]

$$w_{JA}(X_t;\delta_{JA}) = \mu_0 + b'x_t + \tfrac{1}{2}x_t'Cx_t + \sum_{\alpha=1}^{A}\left(\mu_{0\alpha} + 2\sum_{j=1}^{J}\left\{\mu_{j\alpha}\cos(j\kappa_\alpha' x_t) - \nu_{j\alpha}\sin(j\kappa_\alpha' x_t)\right\}\right) \qquad (19)$$
$$= \delta_{JA}'\phi_{JA}(X_t)$$

where x_t is the $K \times 1$ vector consisting of the elements of X_t *excluding* any constant term,

$$C = -\sum_{\alpha=1}^{A}\mu_{0\alpha}\kappa_\alpha\kappa_\alpha', \qquad (20)$$

$$\delta_{JA} = (\mu_0, \mu_{01}, \ldots, \mu_{0A}, b_1, \ldots, b_K, \mu_{11}, \ldots, \mu_{JA}, \nu_{11}, \ldots, \nu_{JA}), \qquad (21)$$

denotes the parameters of the Fourier approximation, and $\phi_{JA}(X_t)$ denotes the vector of corresponding transformations of X_t, including linear and quadratic terms in x_t and the $\cos(j\kappa_\alpha' x_t)$ and $\sin(j\kappa_\alpha' x_t)$ transformations. The κ_α's are $K \times 1$ multiple index vectors used to construct all possible elementary combinations of the explanatory variables (i.e., the x_t) and their multiples. For example, as Chen and Randall [3] note, the typical κ_α's when $K = 3$ would include (1,0,0), (0,1,0), (0,0,1), (1,1,0), (1,0,1), (0,1,1), (1,-1,0), (1,0,-1), etc. The number of these multiple indices (*A*) and the number of multiples (*J*) determines the degree of truncation being used in the Fourier series to approximate $w(X;\beta_S)$. Both *A* and *J*, along with the specific κ_α's to be used, must be selected by the analyst. Some guidance regarding these choices is provided in the literature. For example, Chen and Randall [3] indicate that, in practice, analysts rely on only a subset of the possible multiple indices, excluding those indices that do not "...provide further statistical improvements" [3, p. 331]. As a guide to specific choice of specific indices, Gallant [9] notes that the length of the κ_α's is typically no more than 2 or 3.[7] This would rule out, for example, the multiple index $\kappa_\alpha = (1,-2,1)'$. Finally, Creel [6] observes that, in practice, *J* is usually only 1 or 2. In Chen and Randall's [3] original application, the authors chose *J*=*A*=1, with $\kappa_1 = (1,0,0)$. They note that adding multi-indices or increasing *J* did not significantly increase the likelihood function.

Given the Fourier form approximation to $w(X_i;\beta_S)$, the nonstochastic function $G(X_i;\beta_S)$ in equation (16) is then approximated by:

$$G_{JA}(X_i;\delta_{JA}) = \exp[w_{JA}(X_i;\delta_{JA})]. \tag{22}$$

The second unknown function in modeling CV bid responses is the distribution $\Lambda(\cdot)$. Here the authors rely upon a variant of Gallant and Nychka's [10] semi-nonparametric estimation procedure. The heart of this procedure is the specification of a monotonic transformation of the error term ε_{Si} such that

$$\Gamma[h(u_i)] = \Lambda[u_i], \tag{23}$$

where $\Gamma(\cdot)$ is a known distribution (e.g. the exponential distribution). While the appropriate monotonic transformation function is unknown, Chen and Randall approximate $h(\cdot)$ using the polynomial series:

$$h_r(u) = \gamma_0 + \int_0^u (\gamma_1 + \gamma_2\eta + \cdots \gamma_r\eta^{r-1})^2 d\eta. \tag{24}$$

This structure ensures that the transformation is indeed monotonic.

Substituting the approximations to $G(X_i;\beta_S)$ and $\Lambda(\cdot)$, the log-likelihood function corresponding to the model in equation (16) becomes:

$$\mathsf{L}_S = \sum_{i=1}^{n} no_i \ln\left\{\Gamma\left[h_r\left(\frac{B_i}{G_{JA}(X_i;\delta_{JA})}\right)\right]\right\} + \sum_{i=1}^{n}(1-no_i)\ln\left\{1-\Gamma\left[h_r\left(\frac{B_i}{G_{JA}(X_i;\delta_{JA})}\right)\right]\right\}. \tag{25}$$

One of the advantages of the Chen and Randall [3] estimator is that is can be implemented using standard maximum likelihood routines. Furthermore, the authors prove that if the truncation points used in the two approximations (*JA* and *r*) are increased as the sample size *n* increases, the maximum likelihood estimates of both $\hat{w}_{JA}(X_i;\delta_{JA})$ and $\hat{\Lambda}_r(u_i) \equiv \Gamma[\hat{h}_r(u_i)]$ will converge uniformly and almost surely to the underlying functions $w(X_i;\beta_S)$ and $\Lambda(u_i)$.

The conditional mean WTP is obtained by taking the expected value of equation (16), yielding

$$\mu_X = G_{JA}\left(X_i;\hat{\delta}_{JA}\right)\int_0^\infty \tau\hat{h}_r'(\tau)e^{-\hat{h}_r(\tau)}d\tau. \tag{26}$$

This calculation is performed using numerical integration. The semi-nonparametric model's estimated of the conditional median WTP solves

$$\hat{h}_r\left[\frac{m_X}{G_{JA}\left(X;\hat{\delta}_{JA}\right)}\right] = \ln(2). \tag{27}$$

Due to the nonlinear nature of the problem, a closed form solution for m_X is not readily available. Hence, we solve for median WTP via numerical bisection. In general, the estimated median WTP will not be equivalent to mean WTP. Finally, the conditional dispersion of WTP in the population is given by

$$d_X = \sqrt{G_{JA}\left(X_i;\hat{\delta}_{JA}\right)\int_0^\infty \tau^2\hat{h}_r'(\tau)e^{-\hat{h}_r(\tau)}d\tau - \mu_X^2} \tag{28}$$

A Generalized Maximum Entropy Estimator

Another alternative to standard parametric estimators can be constructed using maximum entropy econometrics.[8] The entropy framework has its roots in information theory and the physical sciences, with Boltzman suggesting as early as the 1870's that *entropy* be used to measure the information content of a distribution. Formally, the entropy index for a discrete distribution is given by:

$$H(p) = -\sum_{j=1}^{m} p_j \ln(p_j). \tag{29}$$

where p_j denotes the probability that the j^{th} event occurs and *m* denotes the total number of possible events. Shannon [26] employed entropy as a measure of uncertainty in communications signals. It was Jaynes [17,18], however, that pioneered the use of the entropy metric as the basis for estimation and inference, particularly for problems that are ill defined or intractable using standard statistical procedures. His *maximum entropy principle* argued for selection of the choice probabilities so as to minimize the information structure imposed on the distribution (i.e., maximize the distribution's entropy) and yet remained consistent with the observed data. Golan, Judge and Miller [11] later generalized the maximum entropy approach to allow for noise in the data, with Golan, Judge, and Perloff [12] adapting the approach to the analysis multinomial response data. It is the Generalized Maximum Entropy (or GME) estimator of Golan, Judge, and Perloff [12] that we adapt to the dichotomous choice CV problem.

In the bivariate discrete choice framework, where the analyst observes either $no_t = 1$ or $no_t = 0$, the maximum entropy (ME) estimator is obtained by solving the problem:

$$\underset{p}{Max}\, H(p) = \sum_{t=1}^{N}\left[p_t \ln(p_t) + (1-p_t)\ln(1-p_t)\right] \tag{30}$$

subject to the K moment conditions:

$$Z'no = Z'p \tag{31}$$

where no is the $N \times 1$ vector whose i^{th} element is no_t, $p_t \equiv \Pr[no_t = 1] = \Lambda(\delta_G' Z_t)$, and Z is the $N \times K$ matrix of covariates assumed to influence the choice probabilities. As several authors note (e.g., [11], [12], and [27]), an interesting feature of the ME estimator is that the resulting first order conditions are *identical* to those obtained when $\Lambda(\cdot)$ is assumed to be the logistic cdf and maximum likelihood procedures are used. Thus, the fitted choice probabilities obtain from the commonly used linear logit model are the same as those obtained using the ME estimator.

The problem with the ME estimator is that it assumes that the moment conditions in equation (31) are non-stochastic. The generalized maximum entropy estimator relaxes this assumption, allowing for an unobserved source of noise and replacing equation (31) with

$$Z'no = Z'p + Z'e = Z'(p+e) \tag{32}$$

where e is the $N \times 1$ unobserved disturbance vector. Following Golan, Judge and Perloff [12], the random disturbance is assumed to have a finite number of support points ($v_t, t = 1,\dots,T$) in the interval [-1,1]. Letting $q_{it} = \Pr(e_i = v_t)$, the noise term can be written in matrix notation as

$$e = Vq = \begin{bmatrix} v' & & & \\ & v' & & \\ & & \ddots & \\ & & & v' \end{bmatrix} \begin{bmatrix} (q_{11},\dots,q_{1M})' \\ (q_{21},\dots,q_{2M})' \\ \vdots \\ (q_{n1},\dots,q_{nM})' \end{bmatrix}. \tag{33}$$

The generalized maximum entropy problem becomes one of choosing both the choice probabilities (i.e., the p_i's) and noise probabilities (i.e., the q_{it}'s) optimally. Formally, this involves solving:

$$\underset{p,q}{Max}\, H(p,q) = -\sum_{i=1}^{n}(p_i \ln(p_i) + (1-p_i)\ln(1-p_i)) - \sum_{i=1}^{n}\sum_{t=1}^{T}[q_{it}\ln(q_{it})], \tag{34}$$

subject to

$$Z'no_i = Z'p + Z'Vq \tag{35}$$

and

$$\sum_{t=1}^{T} q_{it} = 1 \quad \forall i = 1,\dots,N. \tag{36}$$

The above problem involves solving for *n* unknown probabilities (i.e., the p_i's) and *nT* error weights (i.e., the q_{it}'s) using the *K+1* data constraints in equation (35) and the *N* adding-up constraints in equation (36). While the above problem can be solved using standard numerical procedures, Golan, Judge and Miller [11] argue that it is typically easier to solve the equivalent dual problem

$$\underset{\lambda}{Max}\, C = \sum_{i=1}^{n} no_i \lambda' Z_i + \sum_{i=1}^{n} \ln(\Omega_i) + \sum_{i=1}^{n} \ln(\Psi_i) - n\ln(T), \tag{37}$$

where

$$\Omega_i \equiv 1 + e^{-\lambda Z_i} \tag{38}$$

$$\Psi_i \equiv \sum_{t=1}^{T} e^{-v_t \lambda' Z_i} \tag{39}$$

and λ is a $(K+1)\times 1$ vector of parameters. The resulting choice probabilities become

$$p_i = \frac{e^{-\lambda' Z_i}}{1+e^{-\lambda' Z_i}} \tag{40}$$

Equation (40) makes clear the similarity between the GME estimator and the standard linear logit model. In general, the GME estimator is a shrinkage estimator. The structure of the error support (v) will imply how "close" the GME estimator is to the logit estimator asymptotically. As the error support vector widens in coverage of the interval [-1,1], the GME estimates collapse to the origin. As the error support vector narrows around zero, the estimates converge to the ML logit estimates. Allowing v to be wide imposes the most shrinkage on the estimates, which includes the benefit of smaller variance properties. Setting v to be narrow permits the most freedom for the estimates to deviate from zero, however, the cost is less flexibility in the stochastic characterization of the model.

Golan, Judge and Perloff [12] take v to be symmetric about zero with endpoints $\left[-N^{-1/2}, N^{1/2}\right]$. As the sample size gets large, the GME estimator converges to the ML logit estimator. Interestingly, altering the error support vector to be symmetric about zero with endpoints $\left[-N^{-1/2}, N^{1/2}\right]$ implies the GME estimator outperforms the probit model even when the true model is probit (this is shown for a standard normal error distribution by Golan, Judge and Miller [11]). When the true error distribution is not standard and we are forced to estimate the variance, the dominance of the GME estimator wanes. In this case, the GME estimator only marginally outperforms the probit model.

Up until this point, we have reviewed the GME estimator for the bivariate discrete choice problem in general terms. Adapting it to the dichotomous choice CV problem is straightforward. We have that

$$p_i = \Pr[no_i = 1] = \Lambda(\delta_G' Z). \tag{41}$$

A comparison to equation (4), suggests that the analogue to the linear probit model emerges if we set $Z_i = [B_i, -X_i']'$ and $\delta_G = (\xi^{-1}, -\xi^{-1}\beta_G')'$. Using the fact that the choice probabilities in equation (40) have a logistic form, both the conditional mean and median WTP are given by

$$\mu_X = m_X = \beta_G' X_i, \tag{42}$$

with the conditional dispersion of WTP in the targeted population given by

$$d_X = \frac{\pi\xi}{\sqrt{3}}. \tag{43}$$

IV. DESIGN OF THE MONTE CARLO STUDY

The estimators detailed in the previous section provide alternative approaches to analyzing consumer responses to dichotomous choice CV questionnaires. In this section, we describe a Monte Carlo experiment designed to investigate and contrast the performance of these approaches in estimating the characteristics of WTP in a target population, including the mean and median WTP and its dispersion in the population.

The Monte Carlo experiment centers around the construction of an underlying "true" distribution of WTP. We consider four basic distributions: normal, lognormal, uniform and a bimodal distribution. The first two distributions provide settings in which the two parametric approaches (probit and log-probit respectively) provide the correct specifications. The uniform and bimodal distributions were chosen to test more extreme departures from the standard parametric assumptions. The bimodal distribution is constructed as a combination of two standard normal populations, displace from each other by a fixed constant in terms of WTP. This might arise in practice if a significant discrete characteristic of the population (e.g., gender) were excluded from the specification of the nonstochastic portion of the bid function (i.e., $w(X_i;\beta)$).

Table 1 summarizes the four basic distributions considered. The second column provides the equations used to generate observations on WTP_t for each of the "true" distributions. A simple linear form was used for the nonstochastic portion of the bid function. In particular, it was assumed that

$$\begin{aligned} w(X_t;\beta) &= \beta' X_t \\ &= \beta_0 + \beta_1 X_t \end{aligned} \tag{44}$$

where $\beta_1 = 2$ and X_t is a single covariate distributed uniformly on the interval [-30,30]. β_0 was selected for each distribution to insure that the mean WTP was equal to 100. The stochastic component for each of the true distributions was then generated according to the specification in the last column of Table 1. The parameter σ_W^2 measures the dispersion of WTP_t for the typical consumer (i.e., $X_t = 0$) in the population. Formally,

$$\sigma_W^2 = Var(WTP_t | X_t = 0) \tag{45}$$

Four dispersion levels were investigated in the Monte Carlo analysis, with σ_W = 5, 10, 25 and 50. Finally, given observations on WTP_t, simulated survey responses to bid values in a dichotomous choice CV questionnaire (i.e., no_t's) were constructed. In all of the Monte Carlo experiments, we employed a bid design in which the sample was evenly divided into five group, facing bids (i.e., B_t's) of 25, 50, 75, 125 or 175 respectively.

For each of the sixteen possible true distributions (i.e., four distribution types and four dispersion levels), *T=500* samples of size *N=300* were drawn. The four estimators described in the previous section were then applied each sample to estimate the mean, median, and dispersion of WTP in the population. The probit, log-probit, and GME estimators assumed the simple linear form in equation (44) for the bid function. For the Chen and Randall [3] semi-nonparametric estimator, we used the Fourier form:

$$w_1(X_t;\delta) = \delta_0 + \delta_1 \cos(\tilde{X}_t) + \delta_2 \sin(\tilde{X}_t) \tag{46}$$

where $\widetilde{X}_i \equiv 2\pi(X_i + 30)/60$, transforming the covariate X_i to lie in the interval $[0,2\pi]$.

V. MONTE CARLO RESULTS

The primary purpose of CV analysis is typically to characterize the distribution of *WTP* for a specific environmental amenity. Thus, we do not report individual parameters, restricting our attention instead to the performance of the models in terms of estimating the conditional mean and median WTP of the typical observation (i.e., μ_0 and m_0, respectively) and the dispersion of WTP in the sample population (i.e., d_0). Starting with the conditional mean, Table 2 provides a summary of the root mean squared error (RMSE) in estimating μ_0 using the four estimators for each of the sixteen assumed *true* distributions. Bold numbers are used for the lowest RMSE within each distribution.

A number of patterns emerge from Table 2. First, as one might expect, the probit model has the lowest RMSE when the underlying distribution is normal and the log-probit model typically performs best when the underlying distribution is lognormal. What is perhaps more surprising is the generally strong performance of probit for all of the assumed distributions. The probit estimator yields the lowest RMSE for 12 of the 16 specifications. Furthermore, even when probit is outperformed by one of the other estimators, the difference is not substantial. The largest difference emerges when the true distribution is lognormal and $\sigma_W = 25$, with the probit model having a RMSE only 16 percentage points higher than the log-probit model. The GME estimator yields generally similar RMSE's, outperforming probit in one case.[9] The same cannot be said for the log-probit s model. The log-probit model's performance is often substantially worse than that of the simple probit model, particularly when there is sizable dispersion in WTP. When $\sigma_W = 50$, the RMSE for the log-probit model is between 2.5 and 3 times the RMSE for probit. Finally, we note that the semi-nonparametric (SNP) estimator generally does well when the underlying true distribution is relatively smooth. However, when there is a high degree of curvature in the underlying density function (as there is with the bimodal model when $\sigma_W = 50$), the quadratic approximation to $\Lambda(\cdot)$ appears to be insufficient, with a RMSE of nearly nine times that of the simple probit specification.[10]

Table 3 provides a parallel set of results when the focus in on characterizing the condition median WTP (i.e., m_0). The findings here basically mirror those in Table 2. Again, all of the estimators perform well when there is little variability in the underlying population. However, when level of dispersion is high, as is typically the case in actual CV work, the RMSE of the estimated m_0 varies substantially from estimator to estimator. While the performance of the simple probit model is not quite as strong as when we focus on the mean, it still yields the lowest RMSE in 11 of the 16 cases. Again, the gains are the greatest when there is considerable dispersion in the WTP within the targeted population.

Finally, policy makers are often concerned not only with the central tendencies of WTP, but also with its variability or dispersion within a targeted population. Table 4 reports on the ability of the four estimators to characterize the conditional dispersion of WTP (d_0). Surprisingly, the simple probit model excels in this arena as well. Again, in 12 of the 16 specifications, the probit model outperforms both log-probit and the two semi-nonparametric approaches, with a substantially higher RMSE (42% higher) only in the case of the lognormal distribution when $\sigma_W = 50$. The probit model substantially outperforms the other three approaches when the underlying distribution of preferences has a substantial dispersion and is either bimodal or uniform. The RSME of the probit specification is typically 30 to 40 percent of the RMSE obtained by either the log-probit or SNP estimators. While the GME estimator sometimes matches the performance of the probit model, particularly when the level of dispersion is high, the RMSE in estimating d_0 is substantial when the degree of dispersion is small.

The strong performance of the probit specification highlighted in Table 3 through 5 is consistent with earlier comparisons of parametric and nonparametric estimators. Both Horowitz

[15] and Manski and Thompson [23] found that the logit model, similar in nature to probit, typically dominated the more flexible maximum score estimators. Similarly, Huang, Nychka, and Smith [16] found that conventional probit and logit models outperformed cubic smoothing splines. One explanation for the relatively poor performance of the semi-nonparametric estimators is that, by their nature, they rely more heavily upon the data to reveal the shape of the underlying WTP distribution, rather than assumed distributional structures. As Creel and Loomis [7] note, this suggests that they may require both a greater number and range of bid values in order to capture the shape of the underlying WTP distribution. While a full-scale investigation into bid design is beyond the scope of the current paper, Table 5 reports on a simple investigation into the performance of the SNP estimator given a range of bid designs, increasing in complexity from four bid levels to 79 bid levels. The five designs considered place an equal number of bids at various percentiles of the underlying true distribution, with each subsequent bid design essentially doubling the number of bids.[1] As expected, increasing the number and range of the bids does alter the performance of the SNP estimator. However, as in the parametric bid design literature (e.g. Kanninen [19]), the best design for estimating μ_0 differs from the best design for estimating d_0. Estimating the dispersion of WTP benefits substantially from a finer and wider range of bids, whereas estimates of the mean WTP are best with relatively few bid levels.

VI. EMPIRICAL APPLICATION

A common criticism of Monte Carlo studies is that they lack a basis in the real world. Analysts must specify the underlying distributions and functions and choose which characteristics to vary in their experiment. While the hope is always that the choices made bound what one would find in practice, there is always the concern that some critical dimension of the problem

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[1] For this example, we assumed that WTP was normally distributed, with a mean WTP of 250 and a dispersion level of $\sigma_W = 100$.

has been missed.[11] In order to provide additional insight, it is helpful to provide an empirical example. Here, we use the same data base as Chen and Randall [3] employed as an application. The data were obtained from a dichotomous choice CV study design to value improvements to environmental quality of Big Darby Creek in Ohio. The survey was conducted in 1989, yielding information on 274 Ohio residents visiting Battelle-Dargy Creek Park. Table 6 provides a summary of the individual characteristics, while Table 7 provides the pattern of responses obtained in the dichotomous choice CV question. Notice that the survey responses suggest a median WTP of roughly $75, given that 50.9% percent of the population was willing to pay this amount for the water quality improvements. Less than forty percent of the sample was willing to pay $150.

Table 8 provides the mean WTP for the water quality improvements using the four estimators. The probit, log-probit, and GME approaches all yield estimates of the mean *WTP* that lie in the range from $80 to $100. The SNP approach, however, yields a WTP estimate that is roughly four times as large as any of the other approaches. These findings are consistent with the results of the Monte Carlo analysis. In particular, the performance of the SNP estimator is at its worst when there is high level of variability in the underlying distribution of WTP and when the distribution is bimodal, as Chen and Randall [3, p. 334] in their application. The general consistency of the mean WTP estimates when the other three estimators provides some reassurance that the true mean WTP is on the order of $90, it is not precisely measured with any of the models given the limited sample size.

VII. CONCLUSIONS

The purpose of this paper was two-fold, providing an adaptation of the GME estimator to the problem of estimating WTP given dichotomous choice CV data and investigating the relative performance of both parametric and semi-nonparametric estimators using Monte Carlo analysis. One reason for developing and using less parametric approaches is that they, hopefully, limit the role and impact of model specification on the resulting estimates of *WTP*. Our results, however, suggest that nonparametric and semi-nonparametric approaches are not, as yet, a panacea for the problems encountered in using parametric estimators. In fact, the simple linear probit model typically provided the best in estimating the conditional mean and median WTP and its dispersion in the sample, regardless of whether the true distribution of WTP was normal, log-normal, uniform, or bimodal. The GME approach also performed well. However, the log-probit specification, used extensively in the literature to impose non-negativity on the distribution of WTP, did not perform nearly as well. Finally, the SNP estimator did not perform as well when there was substantial curvature in the underlying distribution of WTP. Additional research is needed into this estimator in order to determine how its performance can be enhanced using alternative degrees of truncation in the approximating functions and alternative bid designs.

VIII. REFERENCES

1. L. Adkins, A monte carlo study of a generalized maximum entropy estimator of the binary choice model, *Advances in Econometrics* 12, 183-197 (1997).

2. T. Cameron, A new paradigm for valuing non-market goods using referendum data: Maximum likelihood estimation by censored logistic regression, *Journal of Environmental Economics and Management* 15, 355-379 (1988).

3. H. Chen and A. Randall, Semi-nonparametric estimation of binary response models with an application to natural resource valuation, *Journal of Econometrics* 76, 323-340 (1997).

4. S. Cosslett, Distribution-free maximum likelihood estimator of the binary choice model, *Econometrica* 51, 765-782 (1983).

5. M. Creel, "A Semi-nonparametric, Distribution-free Estimator of Binary Discrete Responses," Revision of Working Paper 267.94, Dept. of Ec. and Ec. Hist., Univ Autonòma de Barcelona (1995).

6. M. Creel, Welfare estimation using the Fourier form: Simulation evidence for the recreation demand case," *The Review of Economics and Statistics*, 79 88-94 (1997).

7. M. Creel and J. Loomis, Semi-nonparametric distribution-free dichotomous choice contingent valuation, *Journal of Environmental Economics and Management* 32, 341-358 (1997).

8. A. R. Gallant, On the bias in flexible functional forms and an essentially unbiased form, *Journal of Econometrics* 15, 211-245 (1981).

9. A. R. Gallant, Unbiased determination of production technologies, *Journal of Econometrics* 20, 285-323 (1982).

10. A. R. Gallant, and D. W. Nychka, Semi-nonparametric maximum likelihood estimation", *Econometrica* 55, 363-390 (1987).

11. A. Golan, G. Judge and D. Miller, "Maximum Entropy Economics: Robust Estimation with Limited Data," John Wiley & Sons, New York (1996).

12. A. Golan, G. Judge and J. Perloff, A maximum entropy approach to recovering information from multinomial response data, *Journal of the American Statistical Association* 91, 841-853 (1996).

13. W. M. Hanemann, Welfare evaluations in contingent valuation experiments with discrete responses, *American Journal of Agricultural Economics* 66, 332-341 (1984).

14. J. Herriges and C. Kling, The performance of nested logit models when welfare estimation is the goal, *American Journal of Agricultural Economics* 79, 792-802 (1997).

15. J. Horowitz, A smoothed maximum score estimator for the binary response model, *Econometrica* **60**, 505-531 (1992).

16. J. C. Huang, D. Nychka, and V. Smith, Discrete choice measures of willingness to pay: A monte carlo evaluation, working paper, September 1996.

17. E. T. Jaynes, Information theory and statistical mechanics, *Physics Review* **106**, 620-30 (1957).

18. E. T. Jaynes, Information theory and statistical mechanics II, *Physics Review* **108**, 171-90 (1957).

19. B. Kanninen, Dichotomous choice contingent valuation, *Land Economics* 69, 1993, 138-146.

20. C. Kling, The importance of functional form in estimation of welfare, *American Journal of Agricultural Economics* **14**, 168-174 (1989).

21. B. Kriström, A nonparametric approach to the estimation of welfare measures in discrete response valuation studies, *Land Economics* 69, 135-139 (1990).

22. C. Manski, The maximum score estimation of the stochastic utility model of choice, *Journal of Econometrics* 3, 205-228 (1975).

23. C. Manski and T. Thompson, Operational characteristics of maximum score estimation, *Journal of Econometrics* **32**, 85-108 (186).

24. R. Matzkin, Non-parametric and distribution-free estimation of the binary threshold crossing and the binary choice models, *Econometrica* **60**, 239-270 (1992).

25. P. Phillips, ERA's: A new approach to small sample theory, *Econometrica* 51, 1505-1527 (1983).

26. C. E. Shannon, A mathematical theory of communication, *Bell System Technical Journal* 27, 379-423 (1948).

27. E. S. Soofi, A generalized formulation of conditional logit with diagnostics, *Journal of the American Statistical Association* 87, 812-16 (1992).

28. T. Stoker, Consistent estimation of scaled coefficients, *Econometrica* 54, 1461-1481 (1986).

29. R. Ziemer, W. Musser, and R. Hill, Recreation demand equations: Functional form and consumer surplus, *American Journal of Agricultural Economics* 62, 136-141 (1980).

Table 1
Monte Carlo Distributions

Distribution	WTP_i	Error Generation
Normal	$WTP_i = \beta_{N0} + \beta_1 X_i + \varepsilon_{Ni}$ $\beta_{N0} = 100$	$\varepsilon_{Ni} \sim N(0, \sigma_W^2)$
Log-normal	$WTP_i = \exp(\beta_{L0} + \beta_1 X_i + \varepsilon_{Li})$ $\beta_{L0} = \ln(100) - \frac{1}{2}\ln\left[1 + (\sigma_W/100)^2\right]$	$\varepsilon_{Li} \sim N\left(0, \ln\left[1 + (\sigma_W/100)^2\right]\right)$
Bimodal	$WTP_i = \beta_{B0} + \beta_1 X_i + \varepsilon_{Li} + \tau_i \Delta$ $\beta_{B0} = 100$ $\tau_i = \begin{cases} 1 & \rho_i > 0.5 \\ -1 & \rho_i \leq 0.5 \end{cases}$	$\varepsilon_{Li} \sim N(0,1)$ $\Delta = \sqrt{\sigma_W^2 - 1}$ $\rho_i \sim Uniform[0,1]$
Uniform	$WTP_i = \beta_{U0} + \beta_1 X_i + \varepsilon_{Ui}$ $\beta_{U0} = 100$	$d_i \sim Uniform[-a, a]$ $a = \frac{1}{2}\sigma_W \sqrt{12}$

Table 2
RMSE in Estimating the Conditional Mean WTP (μ_0)

a. Normal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.5**	**2.0**	3.1	**4.6**
Log-Probit	3.7	5.0	8.4	13.4
SNP	3.3	3.1	5.3	13.8
GME	1.7	2.2	**3.0**	4.7

b. Lognormal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	3.2	3.4	5.0	**4.4**
Log-Probit	**1.3**	**2.1**	**4.3**	11.3
SNP	4.1	3.0	5.4	8.1
GME	4.2	4.1	4.6	4.5

c. Bimodal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.7**	**2.6**	**3.9**	**6.4**
Log-Probit	3.9	5.9	11.3	16.3
SNP	3.2	3.8	13.0	56.5
GME	2.0	2.7	7.6	6.5

d. Uniform Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.4**	**2.1**	**3.2**	**5.2**
Log-Probit	3.6	4.4	8.3	15.0
SNP	3.3	2.9	5.5	18.4
GME	1.7	2.5	3.3	5.4

Table 3
RMSE in Estimating the Conditional Median WTP (m_0)

a. Normal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.5**	**2.0**	3.1	**4.6**
Log-Probit	3.6	4.4	4.9	10.8
SNP	2.6	3.2	4.2	5.9
GME	1.7	2.2	**3.0**	4.7

b. Lognormal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	3.4	3.8	7.5	11.9
Log-Probit	**1.3**	2.0	**4.3**	11.8
SNP	1.6	**2.0**	5.1	**10.4**
GME	4.4	4.5	7.1	11.1

c. Bimodal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.7**	**2.6**	**3.9**	**6.4**
Log-Probit	3.7	4.8	4.8	47.7
SNP	2.7	3.9	21.8	17.0
GME	2.0	2.7	7.6	6.5

d. Uniform Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.4**	**2.1**	**3.2**	**5.2**
Log-Probit	3.5	3.9	5.4	12.2
SNP	2.6	2.9	4.9	8.0
GME	1.7	2.5	3.3	5.4

Table 4
RMSE in Estimating the Conditional Dispersion of WTP (d_0)

a. Normal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.7**	**1.9**	**2.9**	**4.9**
Log-Probit	1.8	2.4	4.7	15.3
SNP	23.8	19.1	6.1	12.6
GME	13.5	10.1	4.4	6.9

b. Lognormal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	1.5	2.3	**3.3**	8.5
Log-Probit	**1.5**	**2.2**	3.4	7.8
SNP	23.6	19.2	5.9	8.4
GME	14.1	10.2	3.5	**6.0**

c. Bimodal Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.9**	**4.1**	**10.2**	**27.2**
Log-Probit	2.5	5.6	16.2	67.1
SNP	23.9	19.5	43.3	67.9
GME	14.0	11.9	80.2	35.5

d. Uniform Distribution

Estimator	$\sigma_W = 5$	$\sigma_W = 10$	$\sigma_W = 25$	$\sigma_W = 50$
Probit	**1.7**	**1.8**	3.1	**6.0**
Log-Probit	1.7	2.2	**2.9**	18.7
SNP	23.8	19.2	6.0	19.5
GME	13.5	10.1	3.5	10.8

Table 5
Sensitivity of SNP Approach to Bid Design

Number of Bids	Design: Percentiles at which bids were evenly spaced	RMSE: μ_0	RMSE: d_0
4	20%,40%,60%,80%	60.2	190.0
9	10%,20%,...,90%	26.3	80.3
19	5%,10%,15%,...,95%	30.3	65.6
39	2.5%,5%,7.5%,...,97.5%	37.8	75.9
79	1.25%,2.5%,3.75%,...,98.75%	33.4	43.9

Table 6
Survey Respondent Characteristics

Variable	Mean	Standard Deviation
Price (B_t)	55.9	50.6
Age	41.4	14.3
Gender	0.50	0.50
Schooling	0.34	0.47

Table 7
Survey Response Patterns

Bid	Percent "No"
$10	28.3
$20	17.6
$30	44.8
$75	49.1
$150	61.5

Table 8
Estimated Mean WTP for Ohio River

Estimator	Estimated Mean WTP
Probit	86.51 (51.15)
Log-Probit	97.66 (91.22)
SNP	391.75 (407.75)
GME	86.44 (52.15)

IX. FOOTNOTES

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[1] It should be noted that valuation efforts based upon revealed preferences (e.g., recreation demand models) are also not immune to the problems of model specification. See, for example, Creel [6], Kling [20], Herriges and Kling [14], and Ziemer et al. [29].

[2] Three notable exceptions are Manksi and Thompson [23]; Horowitz [15]; and Huang, Nychka and Smith [16]. The current paper differs from the first two studies in that [23] and [15] investigate the operational characteristics of the maximum score estimator, which has received little attention in the valuation literature because its implementation can be difficult. In contrast, both of the semi-nonparametric estimators considered in this paper can be implemented using readily available optimization routines. The third study, [16], focuses on the relative performance of the nonparametric cubic smoothing spline, which does not allow for the conditioning of willingness-to-pay on individual characteristics, such as age or income. Both of the semi-nonparametric methods investigated here allow for conditioning variables. Furthermore, Huang, Nychka and Smith start with the specification of an individual's indirect utility function, as in Hanemann [13], whereas we begin by identifying the bid function, as in Cameron [2]. A comparison of results is provided in Section 5 below.

[3] While we will be employing Cameron's [2] bid function approach to analyzing dichotomous choice CV question, parallel results can be obtain when starting with a specification of the individual's indirect utility function, as in Hanemann [13].

[4] The linear logit model is similarly obtained by specifying Λ to be an extreme value distribution. In this case, we would simply replace Φ in the likelihood with the logistic cdf.

[5] Creel and Loomis [7] develop a similar estimator, beginning from a specification of the consumer's indirect utility function (as in Hanemann [13]), rather that starting with bid function.

[6] The notation used in this section is similar to Chen and Randall [3]. Additional details regarding the Fourier form and its characteristics can be found in Chen and Randall [3], Creel [4], and Gallant [8,9].

[7] The length of a κ_α vector in the case of elementary multiple indices corresponds to the sum of the absolute value of its components.

[8] The generalized maximum entropy approach is a relatively recent addition to the econometrics literature. For the sake of brevity, however, this section provides only a brief review of maximum entropy paradigm. A more comprehensive treatment can be found in Golan, Judge and Miller's [11] monograph on entropy econometrics.

[9] This should not be too surprising, given the good fit of the probit model, the well known similarity between the linear probit and linear logit estimators, and the relationship between the GME and logit estimators asymptotically.

[10] This suggests the need for a higher order approximation may be necessary in the case to capture the form of the transformation function $h(\cdot)$. Alternatively, an alternative to the exponential kernal $\Gamma(\cdot)$ may improve the overall fit of the model. However, preliminary investigations along this line did not yield substantial improvements in the RMSE for the SNP estimator.

[11] Analysts will often attempt to minimize this problem by basing the basic model on results obtained previously in the literature. In this case, for example, a mean *WTP* of 100 was chosen to mimic the empirical results obtained in Chen and Randall [3].

Attributes and the Value of a Recreation Experience: A Preliminary Analysis of Wildlife Viewing in Denali National Park

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The events following the grounding of the Exxon Valdez on Bligh Reef in Alaska's Prince William Sound and the accompanying oil spill, in March 1989, have led to an intense scrutiny of the contingent valuation (CV) method of estimating values for nonmarket goods and services. The controversy resulting from the use of CV to establish damages in litigation led to a panel of economic experts, commissioned by the National Oceanographic and Atmospheric Administration (NOAA) of the U.S. Dept. of Commerce to "evaluate the use of CV in determining nonuse values and provide comments to NOAA" (Campbell 1993).

One of the criticisms of CV discussed by the Panel was that "Some of the empirical results produced by CV studies have been alleged to be inconsistent with the assumptions of rational choice" (Arrow et al. 1993). Arrow et al. discuss requirements imposed by rationality:

> Rationality in its weakest form requires certain kinds of consistency among choices made by individuals. For instance, if an individual chooses some purchases at a given set of prices and income, then if some prices fall and there are no other changes, the goods that the individual would now buy would make him or her better off. Similarly, we would expect an individual's preferences over public goods (i.e., bridges, highways, air quality) to reflect the same kind of consistency.
>
> Common notions of rationality impose other requirements which are relevant in different contexts. Usually, though not always, it is reasonable to suppose that more of something regarded as good is better so long as an individual is not satiated. This is in general translated into a willingness to pay somewhat more for more of a good, as judged by the individual. Also, if marginal or incremental willingness to pay for additional amounts does decline with the amount already available, it is usually not reasonable to assume that it declines very abruptly. (Arrow et al. 1993, p. 4604)

Arrow et al. go on to cite some empirical results showing no difference in CV estimated values for what appear to be increased levels of the goods being valued, concluding that such findings make the CV results "hard to explain as the expression of a consistent, rational set of choices." Following from this criticism that CV results are often inconsistent with rational choice, one of the "maladies" enumerated by Arrow et al. that would cause them to judge the findings of a CV study "unreliable" is "inadequate responsiveness to the scope or scale of the environmental insult" (Arrow et al. 1993, p. 4614). This alleged inconsistency with rational choice has become known as the scope problem.

In the years since the NOAA panel, studies have been undertaken attempting to demonstrate that CV can, in fact, produce results that pass a scope test and demonstrate consistency with rational choice. Loomis and Ekstrand (1997), for example, conclude that results obtained using a multiple-bounded CV format pass the scope test when mean willingness to pay for protecting the Mexican spotted owl was compared to that for protecting a group of 62 threatened/endangered species which included the Mexican spotted owl.

The objective of this study was to test whether a link between the value placed on an outdoor recreation experience and the kind/amount/variety of wildlife seen on the trip could be empirically demonstrated. Attempting to empirically demonstrate such a link using an open-ended CV format could also be viewed as a scope test of the CV method. Loomis and Ekstrand used a group of species and an embedded species to test for scope. This study distinguished four levels of a good and estimated values from independent samples.

The Good

The good in this study was a wildlife viewing experience in Denali National Park. The mechanism by which access to the good was offered was a concessionnaire-run tour--the "Tundra Wildlife Tour," a 6 - 7 hour tour into the Park. Informational material for the tour specifically promoted the wildlife that could be seen, and mentioned: "...naturalist will share with you the incredible sights of this diverse land," and "...opportunity to view bear, moose, Dall sheep, and caribou roaming free in their natural habitat." Pictures on the brochure for the Tundra Wildlife Tour were of a caribou, a brown bear, and two pictures of Mt. McKinley. Depending on weather, visibility, and wildlife viewing on the particular day, the tour could go some 75 miles into the Park (the Park road is about 100 miles long). This tour cost $51.00 (half price for children under 12) and tours departed twice each day: early morning, returning mid-afternoon and mid-afternoon, returning late evening.

The tour used school buses (the same as those used for the shuttle service operated by the National Park Service) because the Park road would not accomodate larger, and more comfortable, tour buses.

At the beginning of the Tundra Wildlife Tour, the driver/naturalist talked to the people about what to expect--that viewing wildlife in their natural habitat is highly variable and on any given tour people might see a lot of wildlife or very little. People were told that everyone on the bus needed to help locate wildlife. They were told that whenever they saw wildlife or even thought they saw wildlife they should yell out STOP, and the bus would stop and everyone would have a chance to see whatever wildlife was there. People did this, and it became a very participatory experience. People yelled STOP, then shared binoculars and helped each other spot whatever wildlife anyone found--a lot of wildlife was found by the driver/naturalist as well. Generally, everyone who wanted to saw most or all the wildlife anyone else found or saw. The participatory nature of finding and viewing wildlife, and the sharing of binoculars and helping fellow passengers spot the wildlife made for a more uniform experience for all passengers in terms of the number and kinds of wildlife seen.

Experimental Design

The Tundra Wildlife Tour was selected for several reasons. The Tour specifically promoted wildlife so passengers had an expectation of seeing wildlife. Based on our own experience, and discussions with tour drivers, we knew there was variation in the amount and kind of wildlife seen on any given tour--so different levels of wildlife viewing experiences would occur over a given time period. The tour was run by a private company and a fee was charged, hence it did not seem awkward to ask people how much they would pay to gain access to an experience such as they just had.[1] Perhaps most importantly, this tour was a controlled situation. People got on the bus at the beginning of the tour and off at the end. There were no people joining or leaving the tour enroute. That meant everyone had the same experience and we could collect data on attributes of that experience independent of participant reporting. The National Park Service ran shuttle buses that many visitors used as tour buses (for free at the time of our study). The problem we

[1]In focus groups, during the survey design phase, many people expressed concern over the level of fees charged in a National Park. Comments like: "it should be kept affordable for families" and "the tour should be accessible to everyone" were heard several times. As a result, we tried to remove the context of the National Park as much as possible in an attempt to minimize the effect of such concerns on the results. The final passenger survey contained a cover sheet displaying the State of Alaska seal, and telling participants that the "survey: (1) was being conducted by the State of Alaska; (2) was part of a larger study to collect information about what people think outdoor recreation is worth; (3) would be used to determine whether developing areas <u>outside</u> Denali National Park for tours like this is worthwhile; (4) has <u>nothing</u> to do with setting the price for this tour; (5) has <u>nothing</u> to do with the management of Denali National Park." At the bottom of the cover sheet was a statement saying: "The State of Alaska appreciates the cooperation of the Denali Park Resorts and the National Park Service."

envisioned with using those buses for our study was that people frequently got off one bus at any of several stops and back on some other bus at the same or some other stop. In between times they would sit and view the scenery, or hike around for a few hours, or go off into the backcountry for several days, or anything in between. That meant a variety of experiences with very difficult possibilities for tracking people and collecting data on attributes of their experiences.

Six tour drivers/naturalists participated in the study. Preliminary work on question construction and survey design was followed by an intensive week of focus groups and pretesting, which was then followed by 5 weeks of surveying and data collection. During the period of focus groups and pretesting, researchers rode buses driven by all 6 drivers/naturalists several times taking notes, listening to passengers, talking to drivers, and getting a feel for the tours and the wildlife viewing experience. Three focus groups were conducted (groups were conducted immediately after particular tours with participants recruited from that tour) and several iterations of pretests were carried out on selected tours. During the 5 weeks of data collection, passengers were surveyed on all tours driven by our 6 drivers--a total of 130 bus tours and 4,808 passengers.

Passenger surveys were 7 pages long (page size was half of an 8.5 by 11 sheet) with large print. Surveys were put on clipboards (the same size as the surveys), loaded into duffel bags along with pencils, and put onto the buses. At the end of the tour as the bus was on the way out of the Park, the driver gave a brief introduction to the study and asked people to participate, then distributed the surveys. Surveys were distributed by sliding the duffel bag back along the bus aisle and asking everyone to take a survey and a pencil. Box lunches/snacks and souvenir cups were distributed the same way earlier in the tour so distribution went smoothly. At the drop off points, people dropped the surveys and clipboards into a box at the front of the bus as they disembarked.

Passenger surveys included several ratings of trip satisfaction: whether the tour was worth the price paid (yes/no), overall satisfaction with the tour (10 point scale), satisfaction with the wildlife seen on the tour (10 point scale), and whether the passenger would recommend the tour to a friend (yes/no), along with an open-ended CV question.[2] Passengers were also asked to provide information about tour highlights, tour detractors, some background information about their Alaska trip, and some demographic information. The passenger survey is included as an Appendix.

On each tour, the driver/naturalist kept a log of wildlife seen on the tour along with some other information. The log sheets were designed before the season started with the active participation of some of the drivers and were tested, revised, and retested prior to and during the focus group/pretest period. On the logs, drivers recorded certain information about each "stop," defined as every time the bus stopped to view wildlife.

For each stop the driver recorded the time and location (by odometer reading) of the stop, species seen, number of animals, proximity of the animal(s), ease of seeing the animal(s), activities that were observed, and his or her own subjective rating of overall sighting quality.

[2]On tours conducted by one driver the sample was split, and half answered an open-ended CV question and half answered a multiple-bounded question with opportunites to express uncertainty. For this paper we only analyzed open-ended responses.

Other recorded information included the number of empty seats on the bus, time and point of turnaround, visibility of Mt. McKinley, a general impression of passenger behavior, and any incidents that occurred that might have affected the quality of the experience. Drivers recorded their own subjective rating of overall trip quality, and were encouraged to record anything they thought might help interpret the quality of a particular tour. Drivers were active participants in the whole process. The log sheet is reproduced as Figure 1. Drivers used anywhere from one to six or seven log sheets on individual tours.

Trips were defined as individual bus tours, each containing around 40 people. The "trip data" consist of the information from the driver/naturalist logs. The "passenger data" consist of the information from the passenger surveys.

Analysis and Results

Cluster analysis was used with the trip data (so n=130) to classify trips based on selected trip attributes. The method of clustering used was agglomerative clustering within the SPSS software package. Initially, each case is treated as a separate cluster. From there clusters are grouped based on squared Euclidean distance, so nearest cases are grouped first. Grouping continues until, in the final step, all cases are grouped into one cluster. At each step a distance coefficient is computed. The number of clusters is determined based on rules of thumb and analytical judgement as to when the distance coefficient is "too big," indicating the groups are too far apart to be considered as belonging to the same cluster.

Attributes used for clustering were: (1) the number of "big 5" species (moose, brown bear, caribou, Dall sheep, wolf) seen on the trip, (2) the number of big 5 animals seen on the trip, and (3) the degree to which Mt. McKinley was visible during the trip. Based on those trip attributes, four clusters were identified, with varied amounts of wildlife seen. Trips were clustered independent of any data obtained from passengers. The numbers of trips and passengers in each cluster are shown in Table 1. Trip attributes associated with each cluster (in terms of the cluster variables) are shown in Table 2.

Once trips were assigned to clusters, the passenger data were grouped by the same clusters (based on the particular trip they were on). A comparison of mean satisfaction measures and open-ended CV values by cluster is presented in Table 3. (As shown in Table 2, clusters were numbered in order of increasing numbers of species and animals seen on the trip.) The satisfaction measures track perfectly with increases in viewing success--higher satisfaction is associated with more wildlife seen. Particularly interesting are the CV results. In terms of the highest adult price that passengers were willing to pay, the range is $47.58 for Cluster 1 to $63.49 for Cluster 4. That willingness to pay was converted to a net willingness to pay using the actual $51 tour price (only adults were included in the analysis). In terms of net willingness to pay, the range is -$3.42 for Cluster 1 to $12.49 for Cluster 4.

One-way analysis of variance was performed using cluster as the grouping factor. Results are shown in Table 4. These tests indicate the hypothesis that all group means are equal is rejected at the .05 level for the CV measures and all measures of satisfaction. The ANOVA, however, will reject the null hypothesis if any two cluster group means are unequal, it does not provide information about pairs of clusters. To compare the clusters in more detail, multiple comparison tests were performed using the Bonferroni procedure to test the hypotheses that group means are equal on a group by group basis. The Bonferroni procedure is analogous to doing multiple t-tests between all pairs of groups, but it adjusts the significance level for multiple comparisons involving the same means. The results of the Bonferroni tests indicate that, for the CV values, the null hypothesis of equal means is rejected for all pairs of clusters at the .05 level. For all other satisfaction measures, the null hypothesis of equal means cannot be rejected for clusters 3 and 4, but can be rejected for all other pairs of clusters at the .05 level.

Discussion and Conclusions

Using cluster analysis we were able to differentiate four distinct clusters of wildlife viewing trips within the Tundra Wildlife Tour, independent of any data from passengers. Those clusters represent four levels of wildlife viewing--in effect, four different quality levels of tour, characterized by an increasing amount of the good (wildlife viewing) as one moves from Cluster 1 to Cluster 4. Asking about people's level of satisfaction with the overall tour and the wildlife they saw on their tour led to a conclusion that higher satisfaction was associated with more wildlife seen. Likewise, people expressed a higher willingness to pay, on average, for tours on which more wildlife was seen. This latter finding provides empirical evidence that contingent valuation is capable of picking up positive increments in value for an increased amount of a good--CV results show consistency with assumptions of rational choice to use the terminology of the NOAA panel (Arrow et al. 1993). Results of this study suggest there is nothing inherent in the CV method that leads to results showing inconsistency with rational choice.

References

Arrow, K., R. Solow, P. R. Portney, E. E. Leamer, R. Radner, H. Schuman. 1993. Report of the NOAA panel on contingent valuation. Federal Register, 58(10):4602-4614.

Campbell, T. A. 1993. Natural resource damage assessments under the oil pollution act of 1990. Federal Register, 58(10):4601-4602.

Loomis, J., E. Ekstrand. 1997. Economic benefits of critical habitat for the Mexican spotted owl: A scope test using a multiple-bounded contingent valuation survey. Journal of Agricultural and Resource Economics, 22(2):356-366.

Figure 1. Driver Log Sheet

Date: ____/____/95
Driver Code: ________
Trip Code: ________
Unoccupied Seats: ______

Depart Time:____:____am/pm
Turn-around Time:
____:____am/pm
Turn-around Point: *(circle one)*
Primrose Stony Pt
Toklat ________
Return Time: ____:____am/pm

Incidents on bus that affected the quality of the trip:

Scale: –10 (negative effect) to +10 (positive effect)

rating short description

____ ________________
____ ________________
____ ________________
____ ________________

Mountain Seen:
(circle one)
1. Not Visible
2. Somewhat visible
3. repeated viewings of same portion
4. 1/2 or more
5. 100% visible

Passenger Behavior:
(circle one)
3. Enthusiastic
2. Neutral
1. Contentious

Start Odom. Reading: (at Park HQ)
____

End Odom. Reading: (at Park HQ)
____

Species Codes:
1 = Moose
2 = Bear
3 = Caribou
4 = Sheep
5 = Wolf
6 = Red Fox
7 = Golden Eagle
8 = Ptarmigan
9 = Ground Squirrel
10 = Lynx
11 = Wolverine
12 = Marten
13 = Weasel
14 = Marmot

Proximity Codes:
4 = Close
3 = Moderately Close
2 = Moderately Far Away
1 = Far Away

Ease of Seeing Codes:
4 = Easy to see
3 = Moderately easy to see
2 = Moderately hard to see
1 = hard to see

Activity Codes:
1 = Resting
2 = Feeding
3 = Moving
4 = Other

Overall Codes:
5 = Excellent
4 = Good
3 = Fair
2 = Poor
1 = Very Poor

(The following block is repeated in four columns on the sheet.)

1. Time:
___:___am/pm

2. Odometer reading:________

3. Weather:
(circle one)
a. good
b. fair
c. poor

4. Species:
(write in code)
1 ◯ 2 ◯ 3 ◯

5. # of each Specie:
1 ◯ 2 ◯ 3 ◯

6. Proximity:
(write in code)
1 ◯ 2 ◯ 3 ◯

7. Ease of seeing:
(write in code)
1 ◯ 2 ◯ 3 ◯

8. Activity(ies):
(write in code)
1 ◯ 2 ◯ 3 ◯
1 ◯ 2 ◯ 3 ◯

other: ________
____

9. Overall Sighting Quality: *(write in code)*
1 ◯ 2 ◯ 3 ◯

Table 1

Wildlife Tours

130 buses
4808 participants total

	buses	participants
cluster 1	17	702
cluster 2	48	1884
cluster 3	55	1952
cluster 4	10	270

Table 2

Wildlife Tours - Trip Attributes

	Proportion of tours in cluster	Number of "Big Five" species seen	Number of "Big Five" animals seen	Amount of Mt. McKinley visible
cluster 1*	13%	2.94	5.96	not visible
cluster 2	37%	3.59	10.08	somewhat visible
cluster 3	42%	4.03	14.92	not visible
cluster 4	8%	4.27	21.31	somewhat visible

*the clusters have been numbered in order of increasing wildlife viewing success

Table 3

Wildlife Tours - Satisfaction Measures and Willingness to Pay

	Overall satisfaction*	Satisfaction with wildlife viewing*	Tour was worth the price paid**	Would recommend the tour to a friend**	Highest price WTP	Net WTP (WTP - price)
cluster 1	7.34	6.02	1.69	1.83	$47.58	$-3.42
cluster 2	8.08	7.35	1.79	1.90	$53.45	$ 2.45
cluster 3	8.60	8.24	1.88	1.95	$56.48	$ 5.48
cluster 4	8.69	8.49	1.92	1.96	$63.49	$12.49

* these questions involved a response of a rating on a scale from 1 to 10
** these questions were “yes/no” responses with “yes” coded as 2 and “no” coded as 1

Table 4

Wildlife Tours - Satisfaction and Willingness to Pay - ANOVA

Dependent Variable	F-statistic (d.f. = 3, ∞)
Overall satisfaction	99.753
Satisfaction with wildlife viewing	230.714
Tour price was worth the price paid	56.519
Would recommend the tour to a friend	35.227
Highest price WTP	39.470
Net WTP	39.470

Appendix

The Passenger Survey

IMPORTANT INSTRUCTIONS, PLEASE READ

Because things like the weather, and numbers and kinds of wildlife seen can vary between days and tours, the State needs to learn how people feel about tours under a wide range of conditions. Please answer the following questions based on the conditions that were present for your tour today. Do not consider how you might have felt had conditions been different.

1. The regular price of this tour is $51 (half-price for children under 12). Was the tour you had today worth $51 to you? [WORTHIT]

   1 NO

   2 YES

2. We would like you to consider how much the experience you had today was worth to you so the State can learn about the demand for tours like this one in other areas.

   All things considered - the overall quality of the tour, the scenery and wildlife you saw, the weather - please tell us the highest adult ticket price you could have paid and still felt today's tour was worth it. *(PLEASE FILL IN THE BLANK)* [HIGHPR]

   The highest adult ticket price I could have paid and still felt the tour was worth it is $__________.

3. Please tell us why you answered Question 2 as you did or why you did not answer it. *(PLEASE CIRCLE THE NUMBER OF YOUR RESPONSE)* [WHY]

   1 Because that's what I thought the tour was worth.

   2 I could not put a number on what the tour was worth to me.

   3 Other: *Please explain* _______________________________

   ____[WHYOTHER]

4. Please rate your overall satisfaction with **today's tour** under the conditions **you** experienced. *(PLEASE CIRCLE ONE NUMBER)* [SATTOUR]

Not Satisfied									Very Satisfied
1	2	3	4	5	6	7	8	9	10

5. Now please think about the wildlife you saw (or did not see) during **today's tour.** Consider the numbers and kinds of wildlife you saw, what they were doing, etc., and rate how satisfied you were with the wildlife viewing. *(PLEASE CIRCLE ONE NUMBER)* [SATVIEW]

Not Satisfied									Very Satisfied
1	2	3	4	5	6	7	8	9	10

6. Would you recommend the Tundra Wildlife Tour to a friend? *(CIRCLE ONE NUMBER)* [RECOMM]

1 **NO**

2 **YES**

7. What stood out as the highlight of your tour today?

____

______________________________ [HIGHLIGHT]

8. Were there things that detracted from your enjoyment of the tour? [DETRACT]

1 **NO**

2 **YES** **If yes, what?** ______________

____

______________________________ [WHATDETR]

9. Did you bring binoculars with you today? [BINOC]

1 **NO**

2 **YES**

10. How much of the wildlife pointed out by the bus driver or others did you see? *(PLEASE CIRCLE YOUR RESPONSE)* [HOWMUCH]

1 **All**

2 **Most**

3 **Some**

4 **None**

11. In which U.S. state, Canadian province, or foreign country do you live? ______________ [RESID]

12. **If you do not live in Alaska**, have you been to Alaska before this trip? [TOALASKA]

1 **NO**

2 **YES** **How many times?**

_______ **times before this trip.** [NTOALASK]

13. **Whether you live in Alaska or not**, have you been to Denali National Park before this trip? [TODENALI]

1 **NO**

2 **YES** **How many times?**

_______ **times before this trip.** [NTODENAL]

14. Did you come to Denali as part of a larger tour package? *(PLEASE CIRCLE YOUR RESPONSES)* [TOUPACK]

   1 NO

   2 YES → Was the Tundra Wildlife Tour part of the package you purchased? [INPACKG]

      1 NO

      2 YES → Was the Tundra Wildlife Tour automatically included or was it a separate option you selected? [OPTION]

         1 It was a separate option I selected

         2 It was not a separate option

15. Have you ever taken the Denali Natural Histoy Tour or the Tundra Wildlife Tour before? (*CIRCLE ONE NUMBER*) [BEENTOUR]

   1 NO

   2 YES

16. How many trips (either day or overnight) have you taken on which viewing or photographing wildlife was one of the purposes of the trip? (*CIRCLE ONE NUMBER*) [VIEWTRIP]

   1 None

   2 A few

   3 Several

   4 Many

17. Are you *(PLEASE CIRCLE ONE NUMBER)*: [SEX]

   1 Male   2 Female

18. What is your age? ______ **years old.** [AGE]

19. How many years of schooling have you completed? *(PLEASE CIRCLE ONE NUMBER)* [EDUC]

   1 Eight years or less

   2 Some high school

   3 High school graduate

   4 Some college or technical school

   5 College or technical school graduate

   6 Post graduate work

20. Are you: *(PLEASE CIRCLE ALL THAT APPLY)* [EMPL]

   1 Full-time or part-time student

   2 Retired

   3 Working full-time or part-time outside the home

   4 Self employed

   5 Homemaker

   6 Other

WLO

21. **Including yourself**, how many people live in your household (whether they are traveling with you or not)? (*PLEASE FILL IN EACH BLANK*)

Number:

_____ **People 18 years or older** [HHADULT]

_____ **People under age 18** [HHMINOR]

22. Which of the following categories comes closest to your total 1994 household income? (*CIRCLE ONE NUMBER*) [INCOME]

1 **Less than $25,000**

2 **$25,000 to $34,999**

3 **$35,000 to $49,999**

4 **$50,000 to $74,999**

5 **$75,000 to $99,999**

6 **$100,000 to $149,999**

7 **$150,000 to $199,999**

8 **$200,000 or more**

Thanks for your help!

Table 2

Wildlife Tours - Trip Attributes

	Proportion of tours in cluster	Number of "Big Five" species seen	Number of "Big Five" animals seen	Amount of Mt. McKinley visible
cluster 1*	13%	2.94	5.96	not visible
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*the clusters have been numbered in order of increasing wildlife viewing success

COMPARING WTP ELICITATION PROTOCOLS: A CASE STUDY OF THE VALUE OF AIR QUALITY REGULATION IN WASHINGTON STATE

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ABSTRACT

This paper reports results of a contingent valuation study done as part of an economic appraisal of a proposed state regulation to restrict the burning of bluegrass seed fields. Burning fields enhances yields at low cost to farmers but creates smoke affecting the health and welfare of residents of Eastern Washington. The contingent valuation study used a split payment format with some respondents facing an open-ended question and some asked a DC question. Those asked the dichotomous choice question were also given an open-ended follow up. The three payment formats generated three distinct bid distribution of responses. The results support the contention that different bid formats elicit different bid responses because of informational and strategic differences in the bargaining context.

COMPARING WTP ELICITATION PROTOCOLS: A CASE STUDY OF THE VALUE OF AIR QUALITY REGULATION IN WASHINGTON STATE

In contingent valuation (CV) studies the willingness to pay of the respondents can be elicited through a variety of formats. Which of these elicitation protocols is the best has been a topic of discussion almost since the beginning of the development of the CV method. Periodically the Contingent Valuation community seemingly resolves the issue of the correct bid elicitation protocol, only to enter a new era of debate. Most recently consensus seemed to have settled on the closed-ended, dichotomous choice technique, culminating in its endorsement by the NOAA Panel (Arrow, et al.). However, this apparent consensus gave way to a series of studies--some questioning aspects of the dichotomous choice method, some proposing alternatives such as contingent ranking. So once again the question of correct elicitation format is unsettled.

The issue of the bid elicitation protocol may be merely technical. It may be that different bid elicitation techniques elicit the underlying true willingness to pay (WTP) with different degrees of error and the task is to discover and refine the most accurate method. However, it may be that different bid formats elicit substantively different expressed or stated willingness to pay values. If this latter point is true there is more at stake. In the most negative case generation of different stated WTP distributions may be construed as support for the hypothesis that the CV method is not reliable, that it does not access a true, underlying WTP (e.g., McFadden and Leonard, 1993; Diamond and Hausman, 1993, 1994). Another possibility is that different bid protocols elicit different bid distributions because of differences in the contingent market: the bargaining circumstances. In this case the stated WTP reflects not only the underlying preferences, but also strategic, informational, psychological or other aspects of the bargaining situation (e.g., Ready, et al.; Carson et al. 1997; Carson, 1997). Further investigation is required to distinguish whether different bid elicitation protocols elicit different WTP values, if so, under what conditions, and what components of the stated WTP reflect underlying preferences and which should be attributed to other facets of the protocol.

In this paper we report results from a large-sample survey Contingent Valuation study which included three different bid elicitation formats. Approximately one-third of the respondents faced an open-ended willingness to pay bid format. About two thirds of the respondents faced a closed ended, dichotomous choice bid. In addition, the later group was asked a follow-up open-ended payment question. In initial analysis we have found that responses to the three formats generated three distinct bid distributions. Because the sample was large and the commodity and policy circumstances well defined, initial analysis of the data from this study supports the notion that different payment protocols elicit different bid distributions.

Additional analysis of the data might also provide information useful to understanding the reasons for the discrepancies among bid protocols.

The Study

The study resulted from a proposed regulatory action to improve air quality through restrictions on agricultural burning. Eastern Washington and Northern Idaho are primary centers for the production of Kentucky bluegrass seed. Many of the bluegrass seed fields are located near the most densely populated area (Spokane WA) in an otherwise fairly sparsely inhabited region. In dryland areas bluegrass fields are burned after harvest to remove the residue which covers the crown of the plant. Burning the grass is an economical way to greatly enhances yields in the year following the burn for this perennial crop. However, burning also creates dense plumes of smoke. Epidemiological evidence indicates that smoke is a health hazard to those with respiratory or cardio-vascular conditions like asthma, emphysema or heart disease (Dockery, et al.; Pope, et al.). In fact, the Environmental Protection Agency has recently issued new regulations for airborne particulate matter emphasizing control of the smallest particles, including those generated by combustion processes. The grass field smoke also can be a traffic safety problem, an aesthetic concern, a business and household clean-up cost, and a nuisance interfering with work, recreation or other activities.[1]

After several decades of sometimes contentious discussions, the state air quality agency, the Washington Department of Ecology, proposed a regulation to restrict burning in Washington to two thirds of historic levels. Pursuant to state law, agency officials contracted with researchers at Washington State University to perform a benefit-cost evaluation of the proposed regulation. The Contingent Valuation study reported in this paper was done in the course of that benefit cost evaluation.[2]

At the same time, the researchers were performing a contingent valuation study of potential improvements in air quality from reducing agricultural dust. The Columbia Plateau of Washington, Idaho, and Oregon is subject to large dust events. The Spokane and Tri-cities areas of Washington have among the highest rates of particulate air pollution exceedences (levels which exceed regulatory standards) in the country. A contingent valuation study of the possible air quality benefits from reducing the airborne dust generated from wind erosion in agricultural

[1]Regulations restricting burning of turfgrass fields in Oregon were motivated largely by multi-vehicle freeway accidents caused by restricted visibility from smoke.

[2]The benefit cost study was done under Inter-agency Agreement No C9600164 between Washington State University and the Washington Department of Ecology. The Washington Department of Ecology adopted the regulation in January, 1997. More recently (May 1998) Ecology has adopted a regulation restricting burning on all bluegrass acres with only minor exemptions.

areas was part of a large, multi-disciplinary study sponsored by the Environmental Protection Agency and the Washington State Department of Ecology to address this air quality problem.

In this paper we will report some initial results from the bluegrass field burn restriction study. Similar results were obtained in the dust study. The two air quality studies reinforce each other and provide a useful empirical base upon which to explore contingent valuation procedures. The two studies surveyed overlapping but not identical populations. They had large, random samples of over 1500 respondents each. They involved similar commodities. Furthermore, these commodities are of very direct relevance to the regional households so that issues of existence values and altruism are secondary.

Bid Elicitation--Previous Studies

As noted above, in the last few years the dichotomous choice referendum format has been the preferred payment protocol. The referendum format is believed to be familiar and incentive-compatible (Hoehn and Randall, Mitchell and Carson, Arrow et al.). However, recent studies suggest that the dichotomous choice format can sometimes elicit responses that are higher than the presumed "true" willingness to pay (Kanninen, 1995; Ready et al., 1996;. Herriges and Shogren, 1996; among others). The higher response is due to either anchoring or a kind of strategic response which has been labeled "yea-saying." In yea-saying, the respondent affirms a bid higher than his or her true WTP in order to register approval of the proposed policy. One difficulty in assessing whether dichotomous choice bids are overestimates is that the "true" willingness to pay is not in fact known. In fact, it can be argued that dichotomous choice bids may be strategically under valued if respondents see them as the first in a multi-round "public market" wherein agency officials will lower their offered bid in later rounds--analogous to school bond elections--Randall and Farmer).

Whether or not stated preferences elicited by dichotomous choice payment questions are too high, scholars are amassing considerable evidence that dichotomous choice bids are often higher than open-ended bids (Kealy and Turner, 1993; Boyle, et al, 1996; Ready et al, 1996; and Boyle et al., 1998).[3] Critics of CV cite the evidence of differences in bid values as evidence that the CV method is not only not reliable but that it is not credible (Diamond and Hausman 1993, 1994). However, others argue that the hypothetical public markets used in CV studies have strategic, informational, or other characteristics which may induce systematic and explicable differences among bids (Carson et al., 1996; Ready et al., 1996; Kealy and Turner, 1993;

[3]From a practical perspective, the open ended format has an advantage over the dichotomous choice format in requiring fewer observations. Whereas one only has an upper or lower bound on the respondents value with the dichotomous choice method, one has an observation on the mean value with the open ended format.

Hanemann, 1994; and Carson, 1998). For instance, respondents might view the initial open ended payment mechanism as their initial offer, an offer they would be prepared to pay, but not the highest they would be willing to pay. They would hope to be required to pay less than their maximum WTP. (In contrast, the incentives are to answer the DC question in accordance with one's true maximum WTP if one is convinced that the bid proffered is a single, take it or leave it offer.)

Follow-up bids, whether open-ended or second stage dichotomous choice also appear to have different characteristics than initial bids. Herriges and Shogren (1996) and Boyle et al (1998) have found anchoring of second bids to amounts given in the first bid. Second bids are also likely to have different strategic circumstances since they imply a multi-stage bargaining process instead of a single take it or leave it offer.

Survey Development

Construction of the bluegrass field burn restriction survey instrument started with an instrument developed for the Columbia Plateau agricultural dust study (Scott, et al., 1996). Development of the dust survey began with interviews of key informants about health effects, exposures, and dust generation processes. These interviews and studies in the literature were used to formulate a structured script for two small group focus groups administered by an external professional. This lead to development of a questionnaire which was administered to two panels of about 25 respondents by a marketing organization, Tell-back, Inc. of Spokane, Washington. The panel was non-random and participants were paid a $30. In 90 minute sessions participants responded to questions from a Tell-back moderator using a hand held electronic dialer. The dialer has a continuous scale and feeds directly to a computer to generate a data file. The in-person panel format starts with a predefined questionnaire, but allows discussion and modification of the questions during the session. The dust survey instrument was developed from this information and pretested. For the burn restriction study a new panel questionnaire instrument was developed and administered to two panels in Spokane by Tell-back. The final instrument was pretested by 76 respondents and administered by the Social Survey Research Unit at the University of Idaho. The survey used a random sample from telephone directory data banks using a sample frame provided by Survey Sampling, Inc. of Westport, Connecticut. Advance letters were sent and respondents were placed in a lottery for a small prize ($100) to help increase participation. The study obtained 1561 completed telephone interviews. The response ratio (completes / completes + refusals + did not reach) for the survey was 71 percent. The overall cooperation ratio (completes / completes + refusals) was 77 percent.

Survey Reliability

Before discussing the bid elicitation protocols it is important to consider some general features of the survey instrument that might affect the reliability and validity of the survey results. Analysis of bid elicitation protocols or other specific aspects of a CV study may be possible even if the overall study is not reliable, but certainly one will have more confidence in results from studies in which the underlying commodity (policy) is well formulated and the questionnaire itself is well designed. Overview discussions about the issues involved in constructing a valid, reliable contingent valuation estimate can be found in Arrow, et al. (1993), Mitchell and Carson (1989), and Freeman (1993), and Hanemann (1994) among other sources. Skeptical views can be found in Kahneman and Knetsch (1992) and Diamond and Hausman (1993 and 1994) among others. Discussions by Randall (1998) and Carson (1998) review more recent findings about issues in constructing a valid CV study.

One general issue concerns the nature of the commodity to be valued. Some scholars argue that the contingent valuation method cannot recover valid estimates of existence (passive use) values (see Diamond and Hausman, 1993; 1994; Arrow, 1993) though it may be useful for other kinds of commodities. The commodity in the study, air quality, is a classic public good, but it is one which is used directly. Respondents are valuing real environmental services which affect their health if they are in the at-risk group and their daily life (the risk of traffic accidents, the nuisance of smoke and soot). It may be that there is some component of value that concerns a generic, existence value for "good air," but our preliminary panel studies indicated that direct use components, especially health, are the most important components to area residents.

An extension of the debate over existence values concerns whether altruistic values should be counted in economic studies. Thus, McConnell (1997) argues that altruistic values should only be counted if motivated by "paternalism." In the current study there appears to be an altruistic component to some household values. One of the perceived benefits of clean air for many respondents was its value in reducing risks to others[4]. A follow-up question in the study specifically asked respondents who had no initial stated value, if they would be willing to pay something on the basis of their concern for others. Moreover, in this case there appears to have been altruistic values on both sides of the issue. Both concerns about the health of others and concerns about the impact of the regulations on the well-being of farmers may have affected the expressed WTP. (See discussion below.)

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[4]Thirty percent of respondents agreed that reduction of the health risk to others would be a great benefit and 27% that it would be a moderate benefit to their household. About 36% said that reduction of risk to others would only be a slight or no benefit to their own household. (Missing percentages are unsure/no answer.)

The NOAA panel and others stress that a valid, reliable CV study requires that the commodity (policy) in question be well formulated. In this case respondents were asked to value a very specific policy, one that was being considered for adoption--and indeed was later adopted in precisely the same terms as presented in the survey.[5]

Thus, in terms of a concrete regulation the respondents knew exactly what they were buying. On the other hand, the specific impacts of the regulation were less well defined. The survey included some discussion of the timing, quantity and rationale for bluegrass burning and a general review of the effects of the smoke and the propose reduction. Respondents were also provided statistical information about the likely health effects. But, exactly how much air quality would improve and the extent to which aesthetic, nuisance and health damages would be alleviated could not be clearly described. As discussed below, most respondents were familiar with the general amount of burning and had direct experience of the smoke. Moreover, many public markets (i.e., elections) ask voters to make decisions on even more ill-defined public goods so that respondents are familiar with the task of making decisions when there is some uncertainty about the extent of the effects. Still, the reliability of the study may have been somewhat affected by the lack of precision in the impacts of the regulations.

The NOAA panel and others also direct that respondents be familiar with both the commodity in question and the decision process. In the present study responses reveal that respondents were knowledgeable about local air quality issues generally, and smoke from bluegrass fields in particular. In responses to questions about major contributions to air quality problems respondents answers agreed with local air quality experts. For instance, respondents from Spokane and Northern Idaho rated automobiles and bluegrass smoke as more important contributors to air pollution than did respondents from Eastern Washington who rated agricultural dust more highly (see Table 1). Also, most respondents were personally familiar with smoke from bluegrass (Table 2).

[5] The question read: *The Washington Department of Ecology is currently discussing a law to reduce the burning of field and turf grasses for seed. The intention is to reduce the acres burned in Washington state by 1/3 in 1996 and 2/3 by 1997 and thereafter.* * *The purpose of the rule is to protect the public from the adverse health effects due to smoke.* The asterisk indicates that respondents who wanted more information were read the full text of the proposed registration which included a technical definition of what acres were counted in the base and some exemptions to the restriction.

Table 1: Perceived Sources of Air Pollution (frequencies)

Perceived source	Spokane		E. Washington		No. Idaho	
	Major or Moderate Contributor	**Minor or Insignificant Contributor**	**Major or Moderate Contributor**	**Minor or Insignificant Contributor**	**Major or Moderate Contributor**	**Minor or Insignificant Contributor**
Motor vehicles	**89.7**	9.8	55.5	43.6	**72.6**	26.5
Industrial emissions	70.9	25.6	43.8	55.0	47.9	48.9
Unpaved roads	**75.3**	23.3	55.2	44.3	**69.9**	29.2
Grass burning	67.8	31.2	51.7	48.2	62.6	34.3
Windblown farm dust	66.1	32.0	**71.1**	27.4	50.7	47.0

Note: Figures are percent of responses. Missing percentage is not sure/no answer.

Table 2: Familiarity with Smoke

Question	Response			N
	Never	1-2 times	Over 2 times	
Noticed smoky plume or haze	17%	24%	53%	1467
Experienced dense smoke plume	36%	29%	28%	1217

Besides being familiar, the bid process should also provide a rich decision context. It should be deliberative but neither burdensome nor biased. To provide a decision context and to check understanding respondents were questioned about potential risks of smoke exposure and potential concerns about the proposed regulation. Only a few respondents (about 2%) perceived smoke to be an extreme risk to their household, but about 26% viewed it as a serious or moderate risk. These proportions agree well with the proportion of at-risk households in the sample. Table 3 shows that a little less than 20% of the households had members in the risk group. Also, many households contained people who smoke or who have allergies who might some day enter the at-risk group.

Table 3: Health Status of Household (frequencies)

Health Characteristic	Number in household with characteristic								
	Spokane			E. Washington			N. Idaho		
	0	1	2 or more	0	1	2 or more	0	1	2 or more
Smoke	61.4	16.9	21.6	69.1	13.6	23.3	65.1	18.3	16.6
Allergies	28.4	37.0	34.7	31.1	36.2	32.6	38.7	28.6	37.7
Chronic heart or lung conditions	83.5	13.5	2.9	84.9	12.2	2.9	81.7	16.0	2.3

Note: Numbers are percentage of respondents. Missing percentage is not sure/no answer. N is 1560 for smoke, 1554 for allergies, and 1562 for chronic conditions.

In addition to the health risk, all households might benefit from other aspects of cleaner air. About 10% of households agreed that smoke is a great nuisance, while about 20% deemed it a moderate nuisance. Respondents were not asked about traffic safety, clean-up costs and other possible benefits.

Table 4 shows responses to a set of questions inquiring about concerns that respondents might have with the proposed burn restriction regulation. Responses to these questions show that respondents both understood the negative effects of the restrictions on the agricultural industry and suggest that their responses were influenced by a negative attitude toward the regulation of farmers. (The population of Eastern Washington is generally considered to be politically conservative and supportive of agriculture.) Statistical evidence supports this intuitive understanding of the responses (see below). The pattern of these responses suggest that there was unlikely to have been a large "warm glow" effect. These responses also agree with Portney's proposition (1994) that their may be existence values associated with the non-environmental features of environmental regulations, such as the impact of regulations on the jobs and livelihood of the affected industry. The responses are also consistent with Diamond and Hausman's (1994) assertion that respondents are doing an informal benefit cost analysis of the proposed regulation. But, as Hanemann (1994) observes, informal benefit-cost calculations and other on-the-spot heuristics are part of real world preference formation.[6] We interpret the responses to both the risk questions above and the regulatory concerns questions as evidence of a deliberative process in which households considered both the positive and negative characteristics of the proposed regulations. The responses thereby support the credibility of the results of the survey.

[6] Both Diamond and Hausman and Portney observe that one must be careful that one does not double count both environmental/altruistic existence values and regulatory economic impact existence values.

Table 4: Perceived Concerns about the Burn Restriction Regulation

Concern	Strongly/Somewhat Agree	Strongly/Somewhat Disagree
Rule will cause switch to erosive crops	48.7	27.6
Rule puts financial burden on farmers	80.6	11.6
Wash rule won't work if Idaho farmers burn	66.1	24.9
Rule supporters exaggerate health effects	53.3	37.3
Farmers have right to farm	57.7	37.2
Farmers are being singled out	56.4	38.1
More important community issues (e.g., crime, education)	63.3	20.1

Note: Numbers are percentage of respondents. Missing percentage is not sure/no answer.

One way of testing to see if the commodity is well formulated and more generally if the hypothetical market is valid is a scope test: do respondents vary their willingness to pay with changes in the scope of the commodity. No explicit scope test was built into this survey. However, the study included an implicit scope test. Respondents in different locations are exposed to different levels of smoke. Therefore, a proportionate reduction in smoke would produce lower absolute air quality benefits for someone in a non-bluegrass growing area than from someone located near bluegrass production areas. Indeed, statistical analysis and comments show that respondents from areas with less smoke (Eastern Washington outside Spokane county) were less likely to approve the regulation than those near the smoke (Spokane county and Northern Idaho) (see below). Moreover, one can compare results from the bluegrass study with results from the agricultural dust study. Agricultural dust is more pervasive than bluegrass smoke. Raw WTP values are higher for the agricultural dust study than the bluegrass smoke study--an indirect scope test. Further analysis might be able to quantify the difference in WTP values for different exposures if locations, exposures, and bids can be matched more closely.

Bid Elicitation Protocols

The bid elicitation protocols included a number of features intended to improve the reliability of the bids as well as to facilitate investigation of the influence of different characteristics. One feature was that the willingness to pay question was split into two stages. First, respondents were asked whether they would favor a regulation to reduce burning , and then how much they would pay to see the rule implemented if they approved it. The approval/disapproval question was: ***Suppose you were asked to vote on this smoke reduction***

program reducing the acres farmers can burn by 2/3 of past levels by 1997. Would you vote for or against the program?

The original idea was that, by asking the question in two parts, the incentive to "yea-say" would be reduced. The investigators intended that the respondents contemplate approving the regulation at an implicit zero price. Respondents could express approval for the program at zero cost and would not need to express a symbolic value (if any) in the WTP stage. (Those who did not support the program in this preliminary vote were assigned a zero bid value.) But, Carson (1998b) observed that respondents may well have imagined a positive cost. His observation is supported by a small number of respondents (2) who said whether they would approve or not approve the regulation would depend on the cost[7]. If, indeed, a significant number of respondents inferred their own non-zero cost for the regulation, than the total number of zero values was over-estimated in the study.

Responses to the approval question showed a virtual tie. About 45% of the respondents approved the regulation, 45% opposed it and the rest were either unsure, would not vote, or gave another answer. The fact that the vote was essentially a tie supports the credibility of this study. At least 45% of the respondents did not find the warm glow of supporting clean air sufficient to offer even "cheap" approval.

A logit analysis of responses to this question (table 5) confirms many obvious expectations about those who would favor or oppose the rule. Those respondents who favored the rule had higher incomes and were more likely to be in the at-risk population. They also placed greater importance on the health risks to their own household, the health risks to other households, and the nuisance caused by smoke. Respondents who opposed the rule tended to live in Eastern Washington outside Spokane and felt the rule: singled farmers out, placed financial burdens on farmers, and lacked importance compared to other issues. They also tended to believe that the health risks of clean air are overstated by clean air advocates[8]. This pattern of responses support the notion that respondents undertook a rational, deliberative process in responding to the survey; the preferences appear to be consistent with reasonable expectations. Also, note that respondents less likely to be exposed to smoke were also less likely to support the regulation--supporting the presence of a scope effect.

[7]Unfortunately the investigators failed to anticipate this response even though in hindsight respondents adoption of an inferred price makes sense and can be seen in responses in the focus groups and panel groups. This example underscores that lengthy questionnaire development is only as good as the analysts' ability to use the data.

[8]The possible exaggeration of health risks was also a common theme in the testimony of opponents of the regulation in the public hearings held before adoption of the regulation.

Table 5: Results of Logit Model Predicting Approval/ Disapproval for Rule

Variable	Results
E WA resident	- 0.31 (0.0522)
Health Risk status	0.30 (0.0056)
Risk to others	0.48 (0.0001)
Smoke nuisance?	0.46 (0.0001)
Rule burdens farmers	- 0.39 (0.0001)
Health risks are exaggerated	- 0.31 (0.0001)
Farmers unfairly singled out	- 0.19 (0.0020)
More important issues	- 0.13 (0.0483)
Farmers have right to farm	- 0.26 (0.0001)
Perceived effect, smoke on AQ	0.27 (0.0022)
House income 1995	0.11 (0.0296)

Chi-squared probability values for the test of significance for individual variables are in parenthesis.

Respondents who favored the rule or who were uncertain were asked whether they would pay to have the rule implemented. About 1/3 of the respondents were asked a simple open ended question and about 2/3 were asked a dichotomous choice question: The open ended bid question was:

> *Air Quality agencies want to know how much this reduction program might be worth to your household. Suppose a program were adopted and you and other households had to pay for the program. Program costs would include enforcing the regulations and research into alternatives to grass field burning. Payment would be by some combinations of taxes and fees such as motor vehicle registration, gasoline taxes, property taxes.*
>
> *Now think about how much this program might be worth to your household. Consider how much your household would pay for the benefits of such a program, but remember any amount you pay would leave less money for other expenses. This program is for the reduction of grass smoke only and it would not affect smoke from other sources. How many dollars would your household be willing to pay each year, now and into the future for this program?*

Standard protocols were used to enhance the reliability of the survey. Thus, both the open-ended and dichotomous choice questions were presented in the context of a public market--a vote. The payment question included a budget reminder and substitute/budget competitors reminders. The preliminary questions about concerns also included a substitutes reminder. A set of follow-up/debriefing questions were asked of respondents to help assess the responses.

The questions about of benefits and concerns were presented immediately before the approval/disapproval and bid elicitation questions. If there were order effects in the questions it was intended that they be conservative by presenting the concerns immediately before the payment questions.

One puzzle facing investigators was how to treat the payment vehicle. On the one hand incentive compatibility requires that respondents believe that the payment is not optional. On the other hand this study concerned a regulatory proposal with few credible direct costs. The bulk of costs will be born by the affected industry in reduced rents and, perhaps, by consumers in welfare losses from increased prices. Investigators chose to present a vague tax as the payment vehicle in order to suggest a definite payment but not to invoke a reaction against a tax that would actually not be part of the implemented policy.

Another issue is the choice of the initial values for the dichotomous choice payment question. There is a substantial literature on choice of the initial bid values for dichotomous choice payment questions. Much of the literature concerns the impact of bid choices on precision/efficiency--the nearer that the bid values approximate the distribution of the underlying WTP, the more efficient the dichotomous choice questions (Cooper and Loomis, 1992). However, recent studies suggest that respondents may anchor to bid values, so that appropriate choice of initial bids is crucial (Boyle et al. 1998). In this study bid values were chosen based on the responses given to open ended questions from earlier survey results.[9] Selecting the initial bid values from the presumed distribution of the underlying values should reduce the potential bias created by anchoring.

Open-Ended V. Dichotomous Choice Payment Questions

Initial analysis of the responses to the dichotomous choice (DC) and open-ended (OE) format payment questions reveals very different bid distributions. Figure one shows the open-ended responses (converted to a synthetic dichotomous choice distribution) compared to the (logit) estimated dichotomous choice cumulative percent no bids. (Appendix Table A-2 gives mean values for the combined open-ended bids to show the approximate amounts of these bids.) Visual inspection shows that the probability of a no answer at any given bid level is higher for the

[9]The bid values were based on the open-ended responses of the dust survey. Since values in the dust survey were higher, this may have created an additional upward bias in the grass-burn DC if anchoring was present. But note that both the dust and grass burn studies had open-ended bid distributions below the final DC distribution.

open-ended payment question than the dichotomous choice question. For statistical analysis a simple synthetic dichotomous choice distribution was

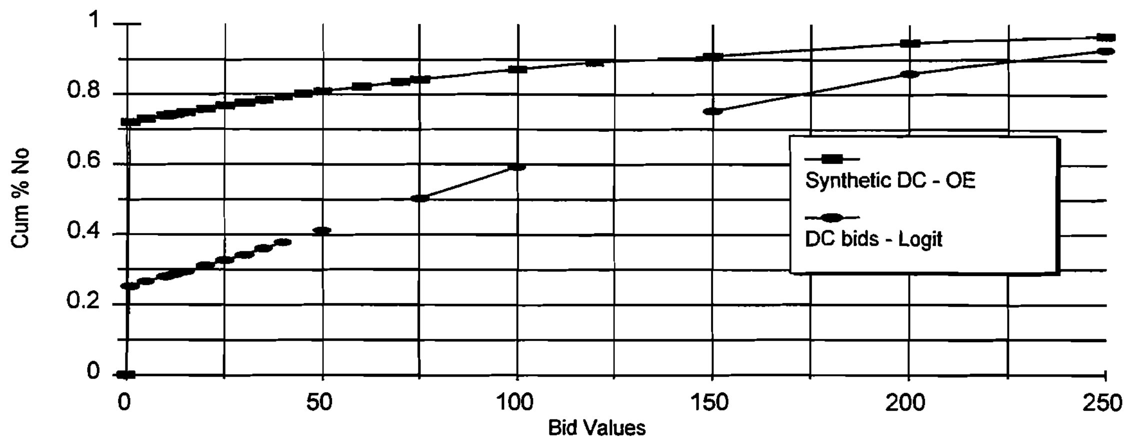

Figure 1. DC v Syn DC(OE)

constructed from the open-ended responses by randomly assigning a bid value to each open-ended observation.[10] Table 6 shows the results of a logit analysis of the dichotomous choice bid distribution compared to the synthetic dichotomous choice distribution inferred from the open-ended responses. The difference between the two distributions is apparent in the radical difference between median bid values. The median bid value for the dichotomous choice question was $74 whereas the median for the synthetic dichotomous choice was $0! In fact, 72% of the synthetic dichotomous choice responses are at zero.

Pooling of the data illustrates the differences statistically. Logit estimates were made for the two distributions, the pooled responses, and for a pooled set of responses with separate intercepts for the two response types. The intercept term for the synthetic responses is clearly significant suggesting that pooling of the data is not supported. While more refined analysis of this data is being pursued, it is clear from even this simple analysis that the distributions are different.

[10]Future analysis will employ Monte Carlo techniques to generate the synthetic distribution.

Table 5: Logit Analysis of DC v Synthetic DC(OE)

	DC	Syn DC(OE)	Combined	Separate Intercept
Intercept	-1.0880	.9441	-.5500	.7498
Pr>chi	.0001	.0015	.0001	.0005
D - Syn				-1.7965
Pr>chi				.0001
Bid	.0146	.0099	.0130	.0139
Pr>chi	.0001	.0233	.0001	.0001
Log Likli			-403.7	-365.6
Median	74	0	42	0
0 threshold	.25	.72	.37	.67
N	463	181	644	644

Initial Open-Ended Versus Follow-up Open-ended Payment Questions

Open-ended questions were also asked as follow-up questions to those who received the dichotomous choice question. Those who said no to the bid offered in the dichotomous choice were asked what the maximum amount they would pay would be, if any. Those who said yes to the offered bid were asked the maximum that they would pay. To date little analysis of the follow-up open-ended distribution has been completed but responses reveal what seems to be still a third distribution (see Figure 2). Visual inspection suggests that at low bid values the follow-up distribution is lower than the initial open-ended but at high bid values the follow-up distribution is higher than the initial open-ended. This results hints at an anchoring effect from the dichotomous choice offered bids. It is also interesting that the follow-up open-ended bids are lower than the dichotomous choice distribution, and, in fact, appear to be lower than the initial open-ended bid.

Discussion

Results of this study confirm earlier studies which found different and higher distributions of bids for the dichotomous choice format than for the open-ended format. The differences were very sharp in this study. Our review of the literature suggested four hypothesis about the meaning and cause of this difference: 1) one or both methods are bad survey protocols and are not measuring the underlying variable; 2) preferences are not stable, since there are no underlying preferences to measure, the WTP values are artifacts of the survey protocols; 3) while

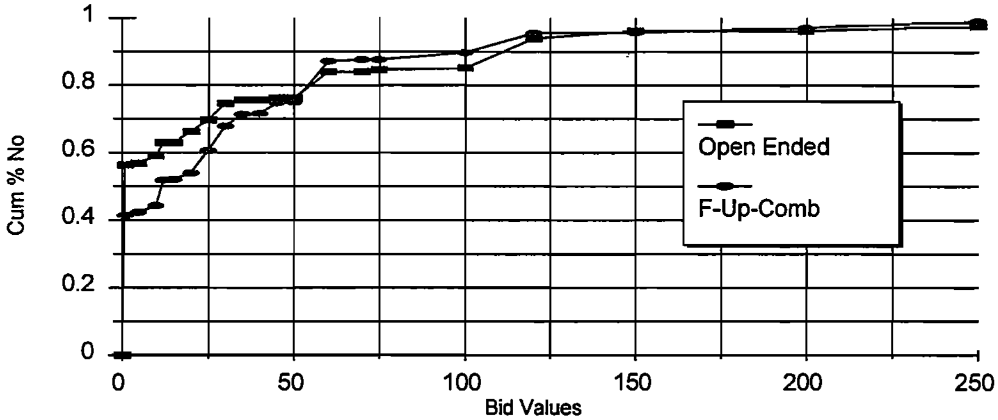

Figure 2. OE v. FU (OE)

underlying preferences are stable, the CV method is intrinsically unable to measure them and again the WTP values are artifacts of the survey protocols[11]; 4)underlying preferences are stable and accessible but different market structures have different informational, strategic, and psychological properties which may lead to differences in the distribution of responses. If either hypothesis two or three is true, one must consider what it is that the CV instrument is measuring. Proponents of these hypotheses often suggest that, if CV studies measure anything, it is a) the undifferentiated "warm glow" of attitudinal support for some valued action rather than a preference for a specific quality or quantity of a public good or b) the result of an informal B/C by respondents about what should be the level of their contribution (e.g., Diamond and Hausman, 1994).

This study contained no definitive result to distinguish among these hypotheses. However, the nature of the commodity itself, the distribution of responses, and indications of a purposive deliberative process and some scope effects discount the "warm glow" explanation. However, the benefit cost version of hypotheses two and three is consistent with the evidence that respondents were influenced by the perceived affect of the regulation on the health of others on the one hand and on the economic status of farmers on the other. Still the authors agree with

[11]This hypothesis is consistent with the revealed preference methodological tradition in economics. The revealed preference methodological position can be traced to the positivist and behaviorist position of the 1930s and 40s which treated all mental processes as subjective, unmeasurable, and unaccessible by scientific procedures. Minds are black boxes and only actual behavior can be the province of science (economics). Economists of this tradition distrust measures of stated preferences and require behavioral links. This tradition also distrusts interdependent utility processes (altruism, malevolence).

Hanemann that incorporation of calculations of benefits and costs and altruism are both consistent with individual utility.

It seems to us the very sharpness of the differences in the DC and OE responses in this case provides the strongest support for hypothesis four. The warm glow and cost-benefit alternative hypotheses do not suggest a mechanism for why these two distributions should be different so consistently across studies and so sharply in this case. If respondents were merely indicating a warm glow than it would seem that the responses to the payment question should be consistent, though imprecise, across bid formats and other features of the survey. Proponents of warm glow suggest that the absence of scope effects supports warm glow because warm glow would not vary with scope. By the same reasoning "warm glow" should also not vary according to the strategic factors of the hypothetical market either.

On the other hand, the differences in strategic and informational characteristics do appear to explain differences between the two bids. Strategic factors suggest that respondents are prepared to pay the OE bid amount, but it is like an initial strategic bid in a bargaining process--respondents hope to receive the commodity for less than the maximum they would be willing to pay--they hope to secure some consumer surplus. If, on the other hand, the WTP bid is constructed to convey a real take it or leave it circumstance, then the strategic incentive is to reveal maximum WTP in the DC bid. The DC bid is therefore an upper bound on what one would expect to obtain from the OE question. The anchoring and possible yea-saying in the DC bid does suggest that these bids may be higher than the true WTP, but it also supports the hypothesis that respondents are responding to the informational and strategic aspects of the contingent market.

Analysis of the comparison between the follow-up open-ended bid distribution and the other two bid distributions is less far along. The shape of the follow-up distribution suggests anchoring to the DC bid amounts. The lower values of the distribution suggest a strategic effect--the second bid is what respondents think would be an acceptable bid; it is not their highest bid.

Conclusions

This study supports studies that show a distinct difference between the dichotomous choice and open-ended bid elicitation formats. It is also consistent with hypotheses that the difference between the two bid distributions may be due to strategic and informational factors. The evidence from this studies and recent study support the proposition that the OE bid in a well formulated CV study is likely to be low relative to true WTP. It also has no evidence to contradict the proposition that the DC bid distribution may be high relative to true WTP because of informational (anchoring) or strategic (yea-saying) effects. These results suggest the need for continued theoretic and empirical investigation of these differences. In the meantime it would seem prudent that studies with policy content use split samples with both OE & DC format until we understand the strategic and informational components of the bid process better. Since the dichotomous choice bids may be high and the open-ended bids may be low, one inference is that decision makers might want to consider them to be informal bounds on true WTP.

Another feature of this study was the use of the two stage referendum to try to anticipate and off-set a putative yea-saying effect. It may be that this method had some effect in reducing yea-saying, but a split sample test would have been necessary to clearly establish an impact. Even if the technique was successful in off-setting yea-saying, the cost was a potential underestimate of the WTP value because some respondents might have voted no under an assumption of a non-zero cost to the program. One thing this experience makes clear is that regulatory studies are likely to have many zero or even negative values.[12] A lesson seems to be that the bid distribution explicitly account for the mass of the distribution at the zero values by including some mechanism for a zero bid.

Another lesson from this study is support for the observation that there may be altruistic and existence values on both sides of an environmental regulatory question. Again, more empirical and theoretic research on this question is warranted.

A final lesson of this study is to look for implicit scope tests. In cases where there is a use component to the public good, different groups may, voluntarily or involuntarily, consume different quantities of the good. In some cases it will be possible to build into studies of this kind of public good a component for a kind of hedonics study of scope.

[12]The study included an attempt to measure these negative values as WTP to avoid the regulation but the answers were few and unstable.

References

Arrow, K., R. Solow, P. Portnoy, E. Leamer, R. Radner, H. Schumar. "Report of the NOAA Panel on Contingent Valuation." Federal Register 58(1993): 4601-14.

Bishop, R. and T. Heberlein. "Measuring Values of Extra-Market Goals: Are Indirect Measures Biased," Am. J. of Agric. Econ., 61(979): 926-930.

Boyle, K.J., F.R. Johnson, D.W. McCollum, W.H. Desvouges, R.W. Dunford, and S.P. Hudson. "Valuing Public Goods: Discrete versus Continuous Contingent Valuation Responses," Land Economics, 72 (August 1996) 381-96.

Boyle, K.J., F.R. Johnson, and D.W. McCollum, "Anchoring and Adjustment in Single-bounded, Contingent Valuation Questions," AJAE. 19(Nov. 1997): 1495-1500.

Boyle, K.J., H.F. Maconald, H. Cheng, and D.W. McCollum. "Bid Design and Yea Sayiing in Single-Bounded, Dichotomous Choice Questions." Land Economics 74(February 1998): 49-64.

Carson, R.T., "Contingent Valuation: Theoretic Advances and Empirical Tests since the NOAA Panel." Amer. J. Agr. Econ. 79(1997): 1501-1507.

Carson, R.T. Personal communication. February, 1998.

Carson, R.T.and R.C. Mitchell. "Measuring Scope in Contingent Valuation." Am. J. of Agric. Econ., (December1993).

Carson, R.T., T.Groves, M.J. Machina. *Proceedings of 1997 Workshop on Valuation and Environmental Policy*. National Science Foundation and Environmental Protection Agency, Arlington VA, 7-8 April, 1997.

Cooper, J. and J. Loomis. "Sensitivity of Willingness to Pay Estimates to Bid Design in Dichotomus Choice Contingent Valuation Models." Land Economics, 68(May 1992): 211-24.

Diamond, P.A. and J.A. Hausman, "On Contingent Valuation Measurement of Nonuse Values," in J. Hausman (ed.) Contingent Valuation: a Critical Assessment. NY: North Holland, 1993.

Diamond, P.A. and J.A. Hausman, "Contingent Valuation: Is Some Number Better than No Number." J. of Economic Perspectives, 8(Fall 1994): 45-64.

Dockery, D.W., C.A. Pope III, X. Xu, J.D. Spengler, J.H. Ware, M.E. Fay, B.G. Ferris, Jr., and F.E. Speizer. 1993. "An Association Between Air Pollution and Mortality in Six U.S. Cities." The New England Journal of Medicine, 329(24):1753-9.

Farmer, M.C. and A. Randall. "Referendum Voting Strategies and Implications for Follow-up Open-ended Responses." in Herriges (ed.) Benefits and Costs Transfer in Natural Resource Planning, Ninth Interim Report of Western Regional Research Project W-133, July 1996.

Freeman, A.M. *The Measurement of Environmental and Resource Values*. Washington, DC: Resources for the Future, 1993.

Hanemann, W.M. "Valuing the Environment through Contingent Valuation," J. of Economic Perspectives, 8(Fall 1994): 19-43.

Herriges, J.A. and J.F. Shogren. "Starting Point Bias in Dichotomus Choice Valuation with Follow-Up Questioning." JEEM. 30(1996): 112-131.

Holland, D., K. Painter, R.D. Scott, P. Wandschneider. "Estimates of the Benefits and costs from Reductions in Grass Seed Field Burning," Dept. of Agricultural Economics, Washington State University, Pullman WA 1997.

Hoehn, J.P. and A. Randall, "A Satisfactory Benfit Cost Indicator from Contingent Valuation." JEEM. 14(1989): 226-47.

Kahneman, D. and J.L. Knetsch, "Valuing Public Goods: the Purchase of Moral Satisfaction," JEEM, 21(1992): 57-70.

Keely, M.J. and R.W. Turner, A Test of the Equality of Closed Ended and Open-Ended Contingent Valuations," AJAE. 75(May 1993): 321-331.

McConnell, K.E. "Does Altruism Undermine Existence Values," J. of Environ. Economics and Management, 32(1997): 22-37.

McFadden, D. and G.K. Leonard. "Issues in the Contingent Valuation of Environmental Goods: Methodologies for Data Collection and Analysis," in J. Hausman (ed.) Contingent Valuation: a Critical Assessment. NY: North Holland, 1993.

Mitchell, R.C. and R.T. Carson. 1989. *Using Surveys to Value Public Goods*. Washington DC: Resources for the Future.

Pope, C.A. III, M.J. Thun, M.M. Namboodiri, D.W. Dockery, J.S. Evans, F.E. Speizer, and C.W. Heath, Jr. 1995. "Particulate Air Pollution as a Predictor of Mortality in a Prospective Study of U.S. Adults." American Journal of Respiratory and Critical Care Medicine, 151:669-674.

Portney, P.R. "The Contingent Valuation Debate: Why Economists Should Care," J. of Economic Perspectives, 8(Fall 1994): 3-17.

Randall, A. "The NOAA Panel Report: A New Beginning or the End of an Era?" AJAE. 79(1997): 1489-1494.

Ready, Buzby, and Hu. "Differences between Continuous and Discrete CV Estimates." Land Economics. 72(August 1996): 397-411.

Scott, R.D., P.R. Wandschneider, and D. Young. Evaluating the Profitability and Social Benefits of Farming Systems from Air Quality Control and Identifying Cost-Effective Air Quality Policies. Annual Report, Profitability and Social Benefits Component; Northwest Columbia Plateau Wind Erosion/PM-10 Project. Pullman Washington, (February 22-23, 1996)

Appendices

Table A-1: Sample by Region

Dispositions	Eastern WA	Spokane	Idaho
Completed interviews	596	746	219
Refusals	133	252	57
TOTAL INELIGIBLE e.g., no listing, moved	287	405	111
Did not reach	71	96	22
TOTAL	1,090	1,500	410
RESPONSE RATE completes/[completes+refusals+did not reach]	74.5%	68.2%	73.5%
COOPERATION RATE completes/[completes+refusals]	81.8%	74.8%	79.4%

Table A-2: Means for Positive WTP Value Responses*

Region	% of Sample	Mean	Std Error
Spokane County (N=246)	33	$49.39	3.49
Eastern WA (N=138)	23	$54.12	4.43
Northern Idaho (N=70)	32	$81.35	18.15

* Responses based on combined split sample open-ended and dichotomous follow-up open-ended questions.

An Approach for Dealing with Uncertain Responses to a Contingent Valuation Question

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1. Introduction

Since the NOAA report on Contingent Valuation (Arrow et al., 1992), researchers have experimented with incorporating "not sure" or "don't know" response categories in dichotomous choice contingent valuation (CV) surveys (Li and Mattson, 1995; Ready et al., 1995; Poe and Welsh, 1996; Wang, 1997). While some practitioners argue against this practice, feeling a "not sure" response category might provide respondents with an easy way out of searching their preferences (see Carson et al., 1995, for a discussion), others deem it important to recognize the degree of respondent confidence about their willingness to pay for a specified environmental good or program.

In some situations, it may indeed be very difficult for an individual to respond to the payment question in a CV survey. Based on our own experience with survey development and analysis, we believe this is likely to happen when individuals are asked payment questions about environmental programs or amenities such that: (i) a large part of the value associated with the program is nonuse value; (ii) respondents are not familiar with the program prior to receiving the CV survey; (iii) the program involves a mix of attributes, some which are perceived to be positive, and some which are perceived to be negative; and (iv) respondents harbor uncertainty about the success of the program.

In addition, dichotomous choice payment questions can sometimes be double-barreled, in the sense that the question captures both the individual's sentiment toward a program and WTP for it. If a respondent favors the program, but feels the offer amount is too high, he or she may have a difficult time responding either yes or no to the WTP question.

In this paper, we analyze the "unsure" responses to a CV question about an environmental program with several of the above mentioned characteristics. Specifically, the survey instrument describes a program to manage noxious weeds in National Forests and asks individuals about their WTP for such a program.

Noxious weeds are non-native plants which tend to dominate other plants within an ecosystem, and in turn decrease the plant and animal diversity within that ecosystem. In addition to threatening the diversity of wilderness areas, noxious weeds may also affect recreational experiences in National Forests.

Controlling noxious weeds can be problematic. Herbicides are reasonably effective, at least in the short term, but there are concerns about the associated environmental risks. Bio-control (e.g., insects) can be used, but there are concerns about introducing a control into an environment where it may not have natural predators. Weeds can also be removed by hand or mowed, and in some cases by letting sheep and cattle graze on them. All of these approaches are controversial. Therefore we expect respondents to feel some ambivalence about the program.

In addition to the standard "Yes" and "No" response categories to the WTP question, we included a "Not Sure" category. We perform a series of analyses to find out who the "unsure" respondents were, why they were unsure and whether we could remedy some of these problems with better survey design. We propose a procedure to help decide whether the "not sure" responses should be interpreted as "no" responses, and statistically model the data based on the outcome of this initial series of tests. We also look at the impact on the estimates of WTP of various methods of dealing with the "unsure" responses.

The remainder of this paper is organized as follows: we describe the survey instrument in section 2. Section 3 describes the data while section 4 considers the determinants of the response to the WTP question.

2. The Survey Instrument

The survey instrument elicits information about willingness to pay for a programs to prevent the spread of noxious weeds, and hence the degradation of existing ecosystems in national forests. The instrument is divided into four sections: (i) a description of National Forests; (ii) a definition of noxious weeds and their impacts on an ecosystem; (iii) a description of the "Noxious Weeds Control Program" and the WTP question, (iv) attitude questions about the environment and demographic questions.

In our preliminary focus groups we discovered participants held erroneous beliefs about national forests. Many mistook national forests for national parks, and had no idea of the spatial distribution of national forests within the United States. To avoid such confusion, a map of the continental United States was placed on the cover page of the survey questionnaire, with national forests colored in green.

The next section introduced Noxious Weeds. Noxious weeds are defined as "undesirable plants that tend to dominate and replace plants in certain areas" (see Box 1). The most important features of noxious weeds, their places of origin and the methods by which they are spread are also described in this section. As shown in Box 1, information is organized in bullets for the sake of brevity and to highlight the most important facts for the respondent.

Focus groups revealed that it was essential for respondents to see pictures of at least some of the most common noxious weeds. Immediately following the basic information box about noxious weeds, we include color photographs of leafy spurge (*Euphorbia Esula*), Canada thistle (*Cirsium Arvense*), and Dalmatian Toadflax (*Linaria genistifolia*), explaining where these plants can be found, where they are originally from, and what kind of damages they cause. Focus group respondents agreed the pictures were vital to the survey. Many people who said they did not know about noxious weeds, recognized the plants in the pictures.

Box 1. Information about noxious weeds in the survey.

<u>What Are Noxious Weeds?</u>

Noxious weeds are undesirable plants that tend to dominate and replace other plants in certain areas.

The term "Noxious" is actually a legal classification for weeds that are so harmful to the environment that federal and state laws require they be controlled by landowners or local government agencies (such as cities and counties).

Noxious weeds:

- were introduced from other areas or countries either intentionally or by mistake
- did not evolve with other plants in these areas
- have no local predators or other enemies such as insects or diseases to control their population growth
- spread quickly once established, taking over native vegetation and causing a decrease in the number of different native species in an area, or even extinction.

Many different plant and animal species live in National Forests. The interaction among these species creates a balance which ensures a healthy forest environment. *Noxious weeds* can upset that balance by crowding out native plants and animals and causing environmental problems.

Based on our focus groups, we felt that the survey instrument needed to emphasize the urgency of the noxious weeds problem, lest respondents fail to understand the need for programs that would monitor and keep under control the spread of noxious weeds. To do so, we present two variants of the same map of the Pacific Northwest, both drawn from BLM documents, in an effort to show the "before" and "after" of noxious weeds. The first variant of the map highlights the (relatively limited) areas where leafy spurge, a common noxious weed, could be found in 1920. The second variant of the map, placed immediately to the right of the first variant, shows how extensively leafy spurge has spread by 1995. Respondents are reminded that "on western public lands it is estimated that the amount of land covered by noxious weeds is increasing by 4600 acres per day. That is about 4600 football fields per day."

After listing options available to control noxious weeds (such as pulling the weeds by hand, burning them, applying herbicides and introducing microorganisms, insects or cattle that feed on the weeds), we finally present the "Noxious Weeds Control Program."

Respondents are told the program – to be administered by the US Forest Service – would be implemented in all national forests in the US. USFS personnel would monitor the spread of weeds and – where necessary – apply the treatment(s) deemed as the most appropriate. Respondents are told applications of herbicides would be calibrated to avoid damages to insects, wildlife, and humans, and to avoid contaminating ground or surface water. Respondents are also told that the program has been tested on some national forests, and has been found successful in 90 percent of the test areas.

Focus group participants were extremely wary of taxes, payments to the government, and timing of their payment as opposed to the timing of the program.[1] This prompted us to specify that the program would be paid for the revenue from a tax imposed on all US households. Revenue from the tax would be placed in an interest-bearing trust fund from which monies would be drawn to run the weed control program for the next 10 years. Reassurance was given that the trust fund would only be used to finance the noxious weeds program.

The payment question followed a dichotomous choice format: if the cost of implementing the noxious weeds program is $X for the respondent's household, in the form of a one-time tax, would the respondent vote in favor or against the program?

The amount $X was varied to the respondent. We assigned an approximately equal number of (potential) respondents to $5, $10, $25, $50 and $75. Respondents were given the option to vote against the program, in favor of the program, or to say that they were unsure. We did not include a follow-up payment question, preferring to use single-bounded data about WTP. We reasoned that, while estimation would benefit from the tighter intervals around the respondent's WTP amount typically obtained with follow-up payment questions, it is often difficult to get respondents to comply with the appropriate branching of the follow-up payments questions. Moreover, Alberini et al. (1997) report that follow-up questions may induce undesirable effects in respondents that alter their WTP for the government provided commodity being valued.

[1] For instance, some participants believed that an on-going, long-term program would require the commitment to pay for many years, which they were not prepared to make. Others objected one-time tax payments for an on-going program. Some participants of them argued that they were in favor of the program, but that the revenue necessary to finance the program should be levied on users of national forests, as opposed to the general public, in the form of an increase in national forest admissions fees. In our survey instrument we decided against this suggestion because it does not allow us to control for the total amount a respondent would be willing to pay.

A section eliciting respondent demographics and environmental preferences concludes the survey.

3. The Data

A total of 110 copies of the survey questionnaire were distributed in late July and early August 1997 to a convenience sample of Colorado residents. The sample is neither random nor representative of the population of Colorado. We received a total of 73 completed surveys. The analysis below refers to this sample of 73 pretest respondents.

Almost all (97.2 percent) of our respondents had visited a national forest. Over 80 percent of the respondents had seen the plants shown in the photographs, but only about 55 percent had heard about noxious weeds and 46 percent were aware that the plants in the photographs are problem plants in some areas. When questioned about the impacts of greatest concern to the respondent, about 72 percent of the sample said that all of soil/water quality, plant and wildlife impacts were important.

Over 60 percent of our respondents were female. Average household income was approximately $60,000 a year, and average age was 39.

Table 1 shows the distribution of responses to the vote question by bid value. About 68 percent voted in favor of the plan and 14 percent voted against, while the "not sure" option was selected by about 18 percent of the sample. The percentage in favor of plan generally decreases as the cost of the plan goes up, whereas both "not sure" and "no" responses would appear to grow with the bid amount.

Table 1. Distribution of Responses to WTP Question by Offer Amount			
Offer Amount	Number of Yes Responses (Percent)	Number of No Responses (Percent)	Number of Unsure Responses (Percent)
$5	14 (29%)	0	1 (8%)
$10	13 (27%)	2 (20%)	0
$25	9 (18%)	2 (20%)	5 (38%)
$50	10 (20%)	2 (20%)	3 (23%)
$75	3 (6%)	4 (40%)	4 (31%)

4. Determinants of Vote Choice and Willingness to Pay

In analyzing the responses to dichotomous choice questions, researchers assume that respondents vote in favor of the plan if their WTP amount is greater than or equal to the amount suggested in the survey, and against the plan if their WTP amount is less than that suggested in the survey. Unfortunately, despite the recommendations of the NOAA panel that a "not sure," "don't know" or "abstain from voting" option be included among the response categories, theory does not provide any guidance as to how such responses should be incorporated into a statistical model and used to estimate the distribution of the underlying WTP variable.

Traditionally, researchers have analyzed responses in the usual way, after reclassifying "not sure" as "no" answers, or after dropping the "not sure" responses from the usable sample. More recently, Wang (1997) has proposed an alternative interpretation. Assuming that respondent answer with definite "yes" or "no" answer only when the bid value is sufficiently far away from the true WTP amount, Wang empirically estimates the bracket around mean WTP within which the bid value must fall for the respondent to choose the "not sure" answer.

In this paper, we illustrate a procedure for diagnosing if and when Wang's approach should be employed. We start with examining whether "not sure" responses should be truly interpreted as distinct from "in-favor" and "against" votes, or should be reclassified as "against" responses. To do so, we fit multinomial logit models explaining the selected answer category as a function of the bid value and other variables that we believe should be related to WTP and/or respondent confidence.

If those respondents who answer "not sure" to the payment question are not different in their observed characteristics from those respondents who answered "no," we re-classify the former as "no" respondents and treat them as such in the standard statistical analysis below. If we find statistically significant differences between "not sure" and "no" respondents, then we must look for alternative statistical models of CV responses. Wang's model is an example of such alternative models.

A. Multinomial logit analyses

A multinomial logit model assumes that an individual selects the response category that ensures the highest utility level. Utility should depend on the bid and other variables capturing taste for environmental quality, prior knowledge of and concern about the problems caused by noxious weeds, income and education. Formally, we express the individual's utility as

(1) $$V_{ij} = x_i\beta_j + \varepsilon_{ij} \qquad (i=1, 2, \ldots, n; j=1, 2, 3)$$

where i indexes the respondent, j indexes the response category, and x contains the bid value, individual characteristics, knowledge and attitudes. The β s are response category-specific coefficients; and the ε's absorb individual and alternative-specific factors known to the individual but not observed by the researcher. If the error terms are independent type I extreme value, it can be shown that the probability of selecting one of the three response categories is equal to:

$$\text{(2)} \qquad \Pr(\text{j is chosen}) = \frac{\exp(x_i\beta_j)}{\sum_{j=1}^{3}\exp(x_i\beta_j)} .$$

Estimation of the likelihood function based on (2) requires that one of the three response categories be treated as the default option and its coefficients be normalized to zero (see Greene, 1997).

In this paper, we treat the "yes" responses as the default alternative. To test whether "not sure" answers are truly distinct from the other vote options, we perform a series of two statistical tests. For each independent variable, the first test is a likelihood ratio test of the null hypothesis that the two coefficients of that variable (the one associated with the "no" response option, and the one associated with the "not sure" response option) are both equal to zero. The second test is a Wald test of the null hypothesis that the two coefficients are equal to one another. If the Wald test does not reject the null hypothesis, we conclude that the "not sure" and "no's" can be combined into a single category.

Results from the multinomial logit estimation are reported in Table 2. We present several specifications. In each specification, we keep the bid value among the independent variables, but change the set of candidate predictors of vote choice. We consider four sets of independent variables: (i) variables measuring the respondent's familiarity with weeds and the problems they cause (AWARE, a dummy indicating whether the respondent has seen the weeds and was aware that they are problem plants); (ii) sociodemographic variables traditionally used in contingent valuation data analyses, such as income, education, gender; (iii) evidence of commitment to environmental organizations (ENVIR, a dummy that takes on a value of one if the respondent contributes money to environmental organizations); and (iv) interest in the environment, combined with proxies for interest in ecosystems and concern about the impacts of the uncontrolled spreading of the weeds. The latter group includes ENVINT (a dummy equal to one if the respondent describes himself or herself strongly or extremely interested in the environment); IMPORT (a variables equal to the sum of the scores circled by the respondent when questioned about how important national forests are as habitat for plants, for fish and wildlife, for their scenic beauty and for wilderness conservation; the higher the score, the more such functions matter to the respondent); and HIIMPACT (a dummy equal to one if all possible noxious weeds impacts are of great concern to the respondent).

The results in Table 2 show clearly that log offer is significantly associated with vote choice: its coefficient is always positive, implying that the higher the offer, the most likely is the respondent to vote against the plan or to report himself or herself as unsure, instead of voting in favor of the plan. The likelihood ratio test (see Table 3) implies that the two coefficients of log offer are statistically significant, while the Wald test implies that they are not significantly different from one another.

The coefficients of most other variables are insignificant -- and in most cases equal to one another. Income has the expected negative sign: the lower the respondent's household income, the more likely is the respondent to vote against the plan or to choose the "not sure" option. However, the two income coefficients are not statistically significant at conventional levels.

Similarly, a higher educational attainment makes more likely for the respondent to vote in favor of the plan, but the effect is not significant.

A stronger degree of interest in the environment makes a person significantly more likely to vote for the plan, as does a higher level of personal concern about ecosystem support functions in national forests. The coefficients of these variables are virtually the same for both "against" voters and "not sure" voters.

To conclude, one possible interpretation for these results is that "not sures" and "nos" are expressing the same preferences – and that "not sure" responses are truly "no" answers, and could be so reclassified.

Table 2. Multinomial logit model. T statistics in parentheses. Omitted category: vote in favor of the plan.

Variable	Specification (i)	Specification (ii)	Specification (iii)	Specification (iv)	Specification (v)
Constant [no]	-5.1597 -3.12	-5.0502 -3.04	-4.2617 (-0.67)	-2.5251 -3.05	-1.1897 -0.39
Constant [not sure]	-4.7688 -3.30	-4.6907 -2.99	1.0014 0.21	-5.4793 -3.36	0.6910 0.25
Log BID [no]	1.1153 2.42	1.0219 2.17	1.4957 2.63	1.1380 2.34	1.3562 2.59
Log BID [not sure]	1.0796 2.64	1.2048 2.62	1.2254 2.69	1.2359 2.83	1.4546 2.79
AWARE [no]		0.5525 0.67			
AWARE [not sure]		-0.8988 -1.17			
Log Income [no]			-0.1218 -0.21		
Log Income [not sure]			-0.3160 -0.68		
MALE [no]			2.3340 2.34		
MALE [not sure]			-0.5584 -0.61		
EDUC [no]			-0.1341 -0.61		
EDUC [not sure]			-0.1644 -0.91		
ENVIR [no]				0.6169 0.79	
ENVIR [not sure]				0.5677 0.82	
ENVINT [no]					-0.4214 -0.54
ENVINT [not sure]					-0.7235 -0.95
HIIMPACT [no]					-1.5108 -1.72
HIIMPACT [not sure]					-1.7777 -2.08
IMPORT [no]					-0.2306 -1.17
IMPORT [not sure]					-0.3498 -1.94

Table 3. Multinomial logit model: tests.

Variable	LR test	Wald test
Log BID	18.1330 (*)	0.4486
AWARE	2.4265	2.2792
Log Income	0.4614	0.0936
MALE	7.2180 (*)	6.3471 (*)
EDUC	1.0079	0.0146
ENVIR	1.0029	0.0032
ENVINT	0.9567	0.0354
HIIMPACT	5.0902	0.0892
IMPORT	3.8557	0.6057

Under the null hypothesis, each LR test is distributed as chi square with two degrees of freedom; the Wald test is distributed as a chi square with one degree of freedom. At the 5 percent significant level, the critical points are 5.99 and 3.84, respectively. (*) denotes statistical significance at the 5 percent level.

B. Analyses of Binary Responses

When the "not sure" responses are reclassified as "no" answers, we can fit a binary response model in which an indicator for the vote choice is the dependent variable, and the bid (or a transformation of it, depending on the distribution WTP is assumed to follow) is one of the independent variables.

We assume that the underlying WTP variable follows the log normal distribution.[2] Hence, we fit to our vote data a probit model in which we enter among the independent variable the logarithmic transformation of the offer. The independent variables are grouped into the same four sets as discussed above. Results are shown in Table 4.

[2] We also fit weibull, exponential, logistic, normal and log logistic distributions. Based on the value of the log likelihood function, we argue that the normal and logistic distributions poorly fit our data, but cannot discriminate between weibull, exponential and log normal distribution, as the log likelihood function is equal to about -23.1 for each of the these distributions.

Table 4. Probit model of vote responses. (Dependent variable equal to one if the vote is in favor of the plan, and zero otherwise. T statistics in parenthesis)

Variable	Specification (i)	Specification (ii)	Specification (iii)	Specification (iv)	Specification (v)
Intercept	2.5230 3.989 (*)	9.0452 0.001 (*)	0.3230 0.129	2.8623 4.015 (*)	-0.4166 -0.282
Log bid	-0.6499 -3.471 (^)	-0.6459 -3.010 (*)	-0.7602 -3.654 (*)	-0.7099 -3.554 (*)	-0.8346 -3.576 (*)
Heard of the plants		-0.9227 -1.899 (^)			
Seen the plants in the pictures		-6.4403 0.002			
Aware that the plants are problem plants		0.7740 1.65 (^)			
Log income			0.1057 0.453		
MALE			-0.3375 -0.888		
EDUC			0.0949 1.067		
ENVIR				-0.3604 -1.038	
IMPORT					0.1833 1.853 (^)
HIIMPACT					0.9846 2.369 (*)
ENVINT					0.3739 1.022

(*) denotes significance at the 5% level or better; (^) denotes significance at the 10% level.

As expected, the coefficient of log bid is negative and strongly significant. The values of the intercept and the coefficient of log bid displayed in column (i) imply that median WTP is equal to $48.54. As always with a log normal distribution, mean WTP is much larger ($158.59).[3]

The coefficient of log bid remains significant in all other specifications. Specification (ii) shows that while WTP may be lower for individuals who have heard about noxious weeds, it is certainly higher for individuals who are aware that the plants in question are problem plants. From specification (iii), it is clear that, holding all else unchanged, WTP grows with respondent income and education, but this effect is not significant. However, the average man's WTP is significantly less (by about 32 percent) than the average woman's. As shown by the coefficient of IMPORT in specification (v), willingness to pay for the noxious weeds plan is positively related to perceived importance of national forests in supporting ecosystems and conservation. The coefficient of HIIMPACT is also positive and strongly significant: respondents are prepared to devote more resources for the plan if they are concerned about all of the possible impacts of noxious weeds we describe in the survey. Finally, persons with strong and extreme interest in the environment tend to have greater WTP amounts, but this effect is not significant.

C. An alternative model for "not sure" responses

Wang (1997) proposes an alternative explanation for "not sure" or "don't know" responses based on the cognitive difficulties faced by respondents when answering the payment question. Wang reasons that respondents select the "yes" responses category only when their underlying WTP amount is sufficiently large relative to the bid value. Similarly, respondents

[3] Cameron and James (1987) show how mean and median WTP can be recovered from the coefficients of the binary

answer "no" to the payment question only when their underlying WTP amount is sufficiently small relative to the bid amount. Respondents whose latent WTP amount is sufficiently close to the bid amount will choose to answer "not sure."

The statistical model of responses corresponding to this argument is that respondents select the "yes" response category when their willingness to pay exceeds the bid plus a certain amount, A. Respondents opt for the "no" response category when WTP is less than the bid, minus a certain amount, A. Finally, undeclared respondents have WTP amounts bracketed within the bid, minus A, and the bid, plus A. Formally, assuming that WTP can be expressed as the sum of a linear combination of individual characteristics and an error term:

$$(3) \qquad \Pr(yes) = \Pr(WTP_i > bid_i + A_i) = \Pr(\frac{\varepsilon_i}{\sigma} > \frac{bid_i}{\sigma} + \frac{A_i}{\sigma} - \frac{x_i\beta}{\sigma}),$$

$$(4) \qquad \Pr(no) = \Pr(WTP_i < bid_i - A_i) = \Pr(\frac{\varepsilon_i}{\sigma} < \frac{bid_i}{\sigma} - \frac{A_i}{\sigma} - \frac{x_i\beta}{\sigma}),$$

and

$$(5) \qquad \Pr(notsure) = \Pr(bid - A_i < WTP_i < bid_i + A_i) =$$

$$= \Pr(\frac{\varepsilon_i}{\sigma} < \frac{bid_i}{\sigma} + \frac{A_i}{\sigma} - \frac{x_i\beta}{\sigma}) - \Pr(\frac{\varepsilon_i}{\sigma} < \frac{bid_i}{\sigma} - \frac{A_i}{\sigma} - \frac{x_i\beta}{\sigma}).$$

Although the results of multinomial logit model and related tests suggest that "not sure" responses could be interpreted as true "no" answers, in this paper we still fit the Wang model (equations (3), (4) and (5)) for the sake of comparing it with the results of our probit models. We estimate the parameters of this model under different assumptions: we first impose that the threshold A is constant for everyone, and then allow it to vary with individual characteristics:

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response model.

$A_i = \gamma_0 + \gamma_1 Z_i$, where Z_i is a gender dummy, log income, educational attainment, and prior knowledge of noxious weeds. The set of regressors x is specified to contain IMPORT, HIIMPACT and ENVINT.

The results, reported in table 5, show that the coefficient estimates are very stable across models, and that only one of 4 possible factors (education, income, gender, and awareness of the weeds problem) is associated with the "threshold" between "not sure" and the other response category. Specifically, the interval around mean WTP that determines a "not sure" responses is much smaller for men.

These models imply a somewhat higher median and mean WTP: \$87.70 and \$296.25, respectively, with A being approximately \$2 in the specification that does not allow A to vary across individuals.

Table 5. Wang model of log WTP. (T statistics in parentheses).

Variable	Specification (i)	Specification (ii)	Specification (iii)	Specification (iv)	Specification (v)
Intercept	2.5309 (2.203)	2.5361 (2.207)	2.6567 (2.371)	2.6291 (2.243)	2.4193 (2.002)
IMPORT	0.0431 (0.573)	0.0426 (0.567)	0.0320 (0.434)	0.0445 (0.578)	0.0534 (0.673)
ENVINT	0.4318 (0.888)	0.4225 (0.865)	0.4127 (0.885)	0.3493 (0.714)	0.4608 (0.909)
HIIMPACT	1.3339 (2.612)	1.3389 (2.611)	1.3315 (2.744)	1.3532 (2.561)	1.3637 (2.525)
σ	1.2089 (3.486)	1.2108 (3.481)	1.1476 (3.557)	1.2018 (3.356)	1.2487 (3.391)
γ_0	0.6289 (2.898)	0.8424 (0.814)	4.0950 (1.431)	0.8747 (2.834)	0.8738 (2.516)
γ_1 -- educ		-0.0137 (-0.214)			
γ_1 -- log income			-0.3234 (-1.240)		
γ_1 -- male				-0.7127 (-2.219)	
γ_1 -- aware2					-0.4344 (-1.252)

D. Nature of "not sure" responses and respondents

To further investigate the interpretation of the "not sure" responses and identify the characteristics of respondents (or of the program) that makes them prone to "not sure" answers, we report in table 6 the distribution of possible reasons for the vote expressed by respondents.

Table 6. Reasons for vote. (Percentage of sample mentioning the listed reason.)

	ALL	IN FAVOR	AGAINST	NOT SURE
Desire to protect the ecosystem	22.6	31.7	0	9.1
The program is worth that much to me	22.6	34.1	0	0
I enjoy nature	1.6	2.4	0	0
National Forests are important	3.2	4.9	0	0
The problem will only get worse	3.2	4.9	0	0
No new taxes, objection to the payment method	11.7	4.9	40.0	9.1
Weeds are a part of nature	3.2	0	40.0	0
Oppose herbicides	4.8	0	10.0	18.2
More info. needed to make decision	1.6	0	0	9.1
Price is too high	14.5	2.4	20.0	54.5
Personally affected by weeds	4.8	7.3	0	0
Other	6.5	7.3	10.0	0

Table 7. Probit analysis. Dependent variable equal to one if 'not sure,' 0 if in favor or against the program. T statistics in parentheses.

Intercept	6.4352 (1.830)
Log income	-0.5413 (1.277)
Male	-0.9583 (1.876)
Educ	-0.0479 (0.451)
Age	0.0538 (1.630)
Envint	-0.6176 (1.407)
Hiimpact	-0.5996 (1.337)
Import	-0.1211 (1.578)

In table 7 we report the results of a probit regression in which the dependent variable takes a value of one if the respondents said he was uncertain on how to vote in the referendum question. The table shows that women and older persons were more likely to declare themselves uncertain about the vote, a result that confirms what was seen in the Wang model. Persons with strong or extreme interest in the environment, respondents who are concerned about all of the possible listed impacts of noxious weeds, and persons who care about the role of national forests in supporting plant and animal life are less likely to be uncertain. Income is also negatively related to the likelihood of being uncertain, but its effect is not significant.

In addition, a separate probit regression that includes log bid as the only independent variable shows that the likelihood of being uncertain about the vote on the program increases with the bid value assigned to the respondent in the survey, once again confirming the implications of the initial multinomial logit model.

5. Conclusions

We have provided a statistical approach for deciding how to best classify unsure respondents. Our approach starts with fitting a multinomial logit model which predicts the probability of selecting a "yes," "no" or "not sure" answer to the payment question as a function of a set of individual characteristics and response category-specific coefficients. We then test for whether the coefficients of regressors associated with "no" and "not sure" answer are equal. If the test rejects the null hypothesis, we proceed with a model formulated by Wang (1997). If the test cannot reject the null hypothesis, we re-interpret the "not sure" responses as "no" responses and fit the usual binary choice models to obtain estimates of WTP.

In our application, we found it makes sense to classify the unsure respondents as "no" respondents. This result is consistent with that of Carson et al. (1995). Results, however, are likely to be specific to each study, survey instrument and program being analyzed. In our application, mean or median WTP does change dramatically with the interpretation given to the "not sure" responses and the statistical model of WTP fit to the data, confirming the importance of the preliminary tests we recommend in this paper.

References

Alberini, Anna, Barbara Kanninen and Richard T. Carson (1997), "Modeling Response Incentive Effects in Dichotomous Choice Contingent Valuation Data," *Land Economics,* 73(3), 309-324.

Alberini, Anna, Kevin Boyle and Michael P. Welsh (1997), "Using Multiple-Bounded Questions to Incorporate Preference Uncertainty in Non-market Valuation," in *W-133, Benefit and Cost Transfer in Natural Resource Planning, 9th Interim Report*, Reno: University of Nevada, Reno.

Arrow, Kenneth, Robert Solow, Paul R. Portney, Edward E. Leamer, Roy Radner and Howard Schuman (1993), "Report of the NOAA Panel on Contingent Valuation," *Federal Register*, 58 (10), 4601-4614.

Carson, Richard T., W. Michael Hanemann, Raymond J. Kopp, Jon A. Krosnick, Robert C. Mitchell, Stanley Presser, Paul A. Ruud, and Kerry Smith, with Michael Conaway and Kerry Martin (1995), "Referendum Design and Contingent Valuation: The NOAA Panel's No-Vote Recommendation," Resources for the Future Discussion Paper QE96-05, Washington, DC (November).

Greene, William H. (1997), *Econometric Analysis*, 3rd edition, Upper Saddle River, NJ: Prentice Hall.

Li, Chuan-Zhong and Leif Mattson (1995), "Discrete Choice under Preference Uncertainty: An Improved Structural Model for Contingent Valuation," *Journal of Environmental Economics and Management*, 28, 256-269.

Poe, Gregory L. and Michael P. Welsh (1996), "Uncertainty in Contingent Values and Procedural Variance Across Elicitation Formats: A Multiple Bounded Discrete Choice Approach," Cornell University Working Paper Series, 96-03, Cornell University, June.

Ready, Richard C., John C. Whitehead and Glenn C. Blomquist (1995), "Contingent Valuation when Respondents are Ambivalent," *Journal of Environmental Economics and Management*, 29, 181-196.

Wang, Hua (1997), "Treatment of 'Don't Know' Responses in Contingent Valuation Surveys: A Random Valuation Model," *Journal of Environmental Economics and Management*, 32, 219-232.

Protest Bids: An Analysis of Respondents to a Survey Using Contingent Valuation

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"Methinks Thou Doth Protest too Much."
Shakespeare

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Abstract

Protest bids in survey research present potential problems when measuring willingness to pay. They are particularly problematic when estimating an aggregate value based on a study of a sample of a population. In this paper we analyze the characteristics of three groups of respondents; those who submitted protest zero bids, those who submitted true zero bids, and those who submitted bids of positive value. We use data from a CV survey of willingness to pay for groundwater protection from nitrate contamination in Southeastern Pennsylvania.

The respondents submitting zero bids were divided into protest and non-protest (or true) bidders based on their answer to a standard follow-up question following the bid elicitation question. Comparisons among the groups were made using logit and multinomial logit regression techniques. Several respondent characteristics were found to decrease the likelihood of submitting a protest bid. Information about water quality and safety, whether it was formal or informal information, not only decreased the likelihood of protesting, it increased the likelihood of offering a positive bid. Respondents with more confidence in the effectiveness of the program in increasing water safety were also less likely to protest and to offer a positive bid. Respondents who have lower perceptions of water safety were less likely to protest, but their willingness to bid did not translate into positive bids.

The only characteristic of respondents that increased the likelihood of protesting was age. Older respondents were more likely to protest but neither more or less likely to bid zero. Gender played no role in either determining protests or positive bidders. Higher income increases the likelihood of offering a positive bid but does not distinguish protest zero bidders from those who submitted non-protest zero bids.

The multinomial regression model showed that few differences exist between the characteristics of respondents who submit protest and true zero bids, but that several distinct differences distinguish protest bidders from positive bidders. In general, the inability to differentiate respondents submitting true zero bids from those submitting protest zeros means that

in this study, protesting some aspect of the scenario or hypothetical plan presented in the questionnaire was not systematically related to social or economic characteristics of the respondent. Tailoring survey design to reduce the number of protest bids may be difficult. On the other hand, differences between bidders submitting protest bids and those submitting positive bids suggest that careful survey design is an essential element of WTP studies.

1. Introduction

The problem of interpreting zero bids in contingent valuation studies is often brushed over as relatively unimportant. A distinction is often made, however, between true zeros--respondents who place a zero value on the good, and protest zeros--respondents who bid zero because they reject some aspect of the questionnaire, not because they actually place a zero value on the good. To combine responses that reflect these different messages introduces bias into the estimates of people's valuation of the good or program being studied. The most common treatment of protest bids in contingent valuation studies is to exclude so called protest bids from the analysis since protest bids do not measure respondents' consideration of the good to be valued.

Making the distinction between true zeros and protest zeros is difficult and prone to error because the reasons underlying zero bids are complex and difficult to elicit through standard survey techniques. Problems may arise if there are systematic differences between those who answer the bid question and those who do not answer, or if differences exist between those bidding a true zero and those offering protest zeros. If contingent valuation surveys are to produce valuation estimates that represent the relevant populations, differences between those who respond and those who do not must be considered. If willingness to pay (WTP) is camouflaged by protests, that is if too many respondents protest, then estimates lose credibility.

Succinctly, protests occur whenever respondents either fail to respond to the bid question, or bid zero for a good which they actually value (Halstead et al. 1992). With mailed surveys, the response rate of the surveyed population is an important determinant of the validity of all theoretical constructs. Non response bias stems from a natural rate of unreturned surveys (due to address changes, deaths, etc.), and protest zero bids (protests) resulting from a respondent's unwillingness to accept the payment vehicle (the method for which the environmental good will be paid for: taxes, donations, etc) or some other aspect of the survey. Correcting for protest bids is more complex than for unreturned questionnaires and can lead to biased estimates because developing an independent determination that a respondent actually values something is subjective. This paper addresses whether differences between respondents who bid legitimately and protests exist, and if they do, what implications they have for demand analysis and ultimately policy.

The issue is considered in the context of policies to protect groundwater from contamination with nitrates. Since the theoretical construct to be measured is willingness to pay for groundwater protection, a respondent who considers the program and registers a zero because he or she believes the program is either unnecessary or not worth the cost is considered a true zero. In this study, those who respond that they do not believe the program will be effective, are not comfortable with the idea of paying into a special fund to protect the environment, or need more information are indicating zero because they do not agree with some aspect of the program and are classified as protest zero. Their WTP may be zero or it may be positive, but from the information provided the researcher cannot tell which. Including protest respondents in the analysis is expected to introduce noise about values not related to true WTP.

Current economics literature on protests is limited. Although Mitchell and Carson (1987), Loomis (1996), Hausman (1992) and Halstead et al. (1992) consider protests, only Halstead et al. are thorough. Mitchell and Carson usually refer to treatment of protests in describing only whether protests were included in mean and median calculations of WTP. Loomis lays out the survey question and defines protests by response to that question of why respondents bid zero but did not intend to provide extensive analysis. In Hausman's (1992) critique of CVM the discussion of protests is scant and inadequate relative to the importance of protests in estimation.

2. Data and Methods

The data for the study comes from research reported in the Master of Science thesis entitled, "Valuing the Benefits of Protecting Groundwater from Nitrate Contamination in Southeastern Pennsylvania" (Delavan, 1997). The study used the contingent valuation method (CVM) to estimate the WTP for groundwater protection and its determinants. A twenty-page questionnaire was mailed to 1000 people in Southeastern Pennsylvania on June 27, 1996. A follow up was sent three weeks later and a third mailing was sent one month after the second mailing. Six hundred and seventeen usable responses were received, 101 were returned as bad addresses and seven were returned as refusals. The response rate was calculated as 617 of 899 or 68.6%. This response rate is comparable to that of other surveys of a general population mailed to households chosen from a telephone list.

2.1 Definition of True Zero and Protest Bids

The questionnaire included a bid question and a follow-up question that explores the reasons that respondents bid zero (see Q22 box below). Approximately one fifth (21.6%) of the 318 who bid zero did so because they did not feel the program would be effective. Few (12.1 %) did not believe the program was necessary and even fewer (7.8%) stated that the program was not worth the cost to them. Nearly 20% did not feel comfortable with the idea of paying into a special fund to protect the environment. Although asked to circle only one answer, many respondents (110 or 34.6% of zero bidders) gave multiple responses. These were classified as "other" leading to a large number of "other" answers. These answers were coded either as one of the first four answers or were included in three additional categories (I) dissatisfaction with government or taxes, (ii) protest related to farming, and (iii) other. (The categorization scheme is available from the authors upon request.)

Q22 Did you write $ 0 in question 21 because: (circle one number)

1 You don't believe the program will be effective.

2 You do not believe the program is necessary.

3 You are not comfortable with the idea of paying into a special fund to protect the environment.

4 The program is not worth the cost to you.

5 You need more information.

6 Other reason (specify) ______________________________

Answers 2 and 4 were classified as reasons that indicate groundwater protection has no value to the respondent (true zero bid). Answers 1, 3, and 5 were classified as protest bids, i.e., the zero response reflects a protest against some aspect of the scenario rather indicating the respondent's value of protecting groundwater. Almost one-third of all respondents (32.1%) offered protest bids. The relatively high proportion of protest bids may have been due to the wording of question 22. The question may have provided too many protest options and had no mechanism to force respondents to circle only one response, although they were asked to do so. When respondents gave multiple reasons for bidding zero, they were classified as protests if one protest reason was included.

Many respondents rejected the payment vehicle (taxes), but the gain in realism from using the tax referendum format was believed preferable to the hypothetical nature of any other possible vehicle.

Bids were classified by elicitation type--dichotomous choice open ended (DOE) and informed open ended (IOE)--and response type--positive, true zero or protest zero (Table 1). Also included in Table 1 are the number of item non responses, called "missing."

2.2 Model and Methods

Each of the four models analyzed employs a qualitative or limited dependent variable. Two models employ a logit model with a dichotomous open ended dependent variable. The first model takes protest = 1 and non protests (both true zeros and positive bidders) = 0. The purpose of this model is to identify differences between those who offered a non protest answer and those who protested. This procedure helps detect problems with the questionnaire or the sample. If, for example, protest respondents did not understand the mailed survey, face to face interviews might be a better alternative. If a number of socioeconomic factors differ, the treatment of protest zeros may bias the result. For example, if the income of those who gave non protest answers differs from those who protested, the true willingness to pay may be different from the estimated willingness to pay, depending on whether or not protest zeros were included in the data for estimation.

In the second Logit model the dependent variable has protests and true zeros equal to zero and the positive bidders equal to one. This test determines if a survey is biased towards a certain group of positive bidders. The merit of the survey can be based on the degree to which bias may have changed the results.

The third model employs a multinomial logit model to determine the extent to which the interpretation of the effects of the different independent variables changes relative to the interpretation of the first two models. In this model the dependent variable takes on three values "0" if protest, "1" if a true zero and "2" if the respondent bid positively. The coefficients on the independent variables are interpreted as relative to the zero or protest category.

The fourth model tests how sensitive the results are to the definition of protest bids. Protest bids are re-categorized so that a narrower definition of protest is used. If the definition of protests is too broad, it may skew willingness to pay. For example, respondents may really be offering a true zero even if they report that they do not think the program is necessary or do not believe the program would be effective. In these cases it is possible that the program was not worth the cost to them because it would not be effective or it was not necessary. These categories are combined into one category. The dependent variable in the multinomial regression now becomes $y = 0$ if respondents offered a positive bid, $y = 1$ if respondents did not believe the program would be effective, $y=2$ if respondents indicated the program was not necessary, or that it is not worth the cost, and $y = 3$ for all others (prior protests minus those in category $y = 1$).

2.3 Independent variables

The independent variables for each model are based on economic theory. Appendix A gives a list of the independent variables and their descriptions.

Several hypotheses can be tested. First, individuals with more information are hypothesized to be more likely to offer a non protest bid. Information takes several forms, however, and theory does not suggest how specific information may affect comparisons. Several questions were asked about information available to the respondent. First, respondents were asked if they had tested their water. The variable YTEST (a dummy with y=1 for 'tested within the past five years' and y=0 for 'not tested') was constructed from responses. It was hypothesized that respondents who made the effort to test their water were less likely to protest. Second, respondents were asked if they knew of high nitrate levels in the region (Q11) or knew of the potential health effects of nitrates (Q12). If so, it was expected that they too would be both more likely to offer a non protest bid, and furthermore to bid positively.

Next, an information section provided respondents with information about groundwater and nitrates in groundwater. This section was followed immediately by an information "quiz" which presented multiple choice questions worded verbatim from the information section. From this "quiz" a variable indicating respondent knowledge of these issues was designed (SMART = 1 if respondents answered at least four of six questions correctly, 0 otherwise). It is not necessary to know if respondents with SMART=1 had prior knowledge or read the information section or both. Again, it is expected that the estimated coefficient on SMART will be positive for legitimate bidders, and positive and larger for positive bidders.

The fifth information variable is HIGHER ED, a dummy variable derived from Q30--EDUCATION, where respondents with at least some college education receive a value of 1 and those without any college education receive a 0.

It is expected that AGE, derived from question 29, will increase the likelihood of both protests and true zeros. Hypothesized reasons for this relationship might include the fact that some respondents are tired of a lifetime of paying taxes or because older respondents will have a shorter time horizon to enjoy the possible benefits of the groundwater protection program.

MALE is a dummy variable representing gender (=1 if male, =0 if female). It is expected to be negative based on results of previous studies indicating that men are less likely to support environmental causes than are women.

Long time residents of the area may be more likely to want to protect groundwater because they have a stake in the community. YEARS is a continuous variable representing the number of years the respondent has lived in the community. MOVE represents respondent's expectation of moving in the next five years (0=definitely not, 1=certainly will). It is anticipated that the more likely a respondent is to move, the more likely they will be to protest or offer a zero bid, since they will not receive benefits from the 10 year program, although they would be paying higher taxes if the program were implemented.

Respondents using private wells (PRIVATE WELLS = 1) are expected to be more likely to offer a positive bid relative to those who get their water from municipal supplies. Those receiving municipal water are already protected by regulations that require testing of municipal water supplies, whereas the private well owners do not have that protection and would be subsidized in the protection of their water source if the program were implemented. Private well owners are expected to be less likely to protest, because the proposed program is directly benefiting them. Conversely, municipal water users may be more likely to protest since they are paying for something from which they do not directly benefit.

Two variables represent people's perceptions. The first, DIFFERENCE, measures respondents' subjective perceptions of the effectiveness of the program. DIFFERENCE is calculated by subtracting people's evaluation of the safety of their drinking water without the program from their evaluation of safety with the program. The higher the DIFFERENCE, the greater we expect the respondents propensity to offer a legitimate bid, and greater yet their propensity to bid positively.

In prior research estimating WTP with a Tobit model, DIFFERENCE was consistently one of the most important factors in determining WTP.

The second perceptions variable is H2O-CONCERN, a dummy variable which measures respondents' concern about groundwater. H2O-CONCERN equals 1 for respondents who indicated a high concern for groundwater safety and believed local government should place a high priority on protecting groundwater. Like DIFFERENCE, HIGHER ED and SMART, the coefficient on H2O-CONCERN is expected to be increasing for both legitimate zeros and positive bids.

Infants and young children are the objects of special concern relative to water contamination. Households with young children are expected to be less likely to protest and more likely to offer a positive bid. The variable CHILD-PRESENT represents the presence of children under the age of four years in the household.

3. Regression Results and Analysis

Three variables were significant in the first regression (Table 2) with PROTEST as the dependent variable. More highly educated respondents (HIGHER ED) and respondents who expect greater effectiveness of the program in increasing water safety (DIFFERENCE) were less likely to protest and older respondents (AGE) were more likely to protest. The fact that DIFFERENCE had a positive influence on consideration of the program is not surprising nor was the result that more highly educated respondents were less likely to protest. The estimated coefficient for AGE was also in line with *a priori* expectations. Respondent's perception of safety (SAFETY) was inversely related to the likelihood of protesting, as was hypothesized. Higher values of this independent variable indicate greater safety of the water supply.

The null hypothesis, that other coefficients are equal to zero, cannot be rejected for all other variables. This might be explained by interactions between independent variables or simply that differences do not exist. Inexplicably, the presence of children in the household was insignificant in all regressions, as was obtaining water from a private well. The insignificance of PRIVATE-WELL may be due to the fact that some private well owners perceive water as a private good, and that these well owners do not appreciate government intervention in the management of their water. Gender did not play a role in explaining protest behavior, nor did income.

The second comparison, examining bidders giving positive values versus those giving zero values (true zeros and protest zeros combined) yielded more interesting results which are shown in Table 3. Again, respondent's perceptions of the effectiveness of the program (DIFFERENCE) increased the likelihood of offering a positive bid. The more information a respondent possessed about groundwater and nitrates in groundwater and the higher their income the greater the likelihood of offering a positive bid. Our expectations were supported in that as income increases the likelihood of offering a positive bid increases relative to offering a zero bid. Similarly, perceptions of the effectiveness of the program, and knowledge of groundwater and nitrate contamination increase the chances of a positive bid relative to a zero bid.

The first multinomial regression model showed that few differences exist between protest and true zeros (Table 4). One factor differentiating the two groups is the perception of present safety, with true zeros more likely to rate current safety higher. More interesting are the coefficients relative to positive bidders versus protests. The results are similar to those in Table 3 comparing positive bidders with zeros and protests. This result may be due to the true zero category including too few observations or it may validate the quality of the survey design by showing that people who respond with protests and with true zeros are similar.

Some reviewers of the research observed that the classification of zero responses as true zeros or protest zeros was flawed. They suggested that the screening response, "You don't believe the program will be effective." is not really a protest. Rather, they consider it a reflection of a statement of the economic value of such a program. To examine this suggestion, we reclassified the zero responses by placing those who indicated Response 1 to Q22 (program not effective) in the true zero category. This classification produces what are referred to as "new zeros" in the following. The second multinomial regression model provides a way of comparing positive bids (the base category) with old zeros, new zeros and the restricted protest category (Table 5). The coefficients describing the connection between positives and old zeros was fascinating because it provides a measure of those factors affecting WTP. In Epp and Delavan (1997) a Tobit model was employed that identified factors affecting WTP, but in that model the coefficients and their significance corresponded to a continuous measure of WTP. Here, the comparison shows what factors determine whether or not a respondent will offer a positive bid.

If the new classification of protest zero bids is truly different from the old classification, then we expect that $B1-B2 \neq 0$ and $B2 = B3$. The results reject the hypothesis that new zeros are different from old zeros (Tables 4 and 5).

4. Conclusions

Protest bidding in CVM represents one of the most serious challenges to validity. If respondents are protesting because they cannot understand the questionnaire or because they are reacting psychologically to some aspect of the survey other than the value of the good or service being studied, then results are prone to misinterpretation and liable to legitimate criticism. If questions designed to identify protests are precise and provide mutually exclusive responses, then it is possible to correct for protests in the interpretation of responses. Furthermore, it may be possible to use this information to design better surveys and reduce protest bidding.

This is not an easy task, and minimizing the number of protest responses involves costs that may or may not be justified depending on the good to be valued and research goals. For example, in this study changing the payment vehicle from a tax to an indirect payment vehicle with less emotional connotations may reduce the number of protests, but it may also introduce a less believable scenario or may lead to inflated estimates of willingness to pay. If the goal is to measure the demand for groundwater protection this inflation may lead to less efficient decisions and may reduce confidence in CVM.

In this paper factors were identified which differentiated protest bidders from non protest bidders, and zero bidders (both protests and actual zeros) from positive bidders, in support of a groundwater protection program in Southeastern Pennsylvania. In general, information of different types (formal or informal) decreased the likelihood of protesting and increased the likelihood of offering a positive bid. Respondents with more confidence that the program will increase water safety were also less likely to protest and more likely to offer a positive bid. Older respondents were more likely to protest but neither more or less likely to bid zero. Higher income increased the likelihood of offering a positive bid but does not distinguish protests from non protests. Lower perceptions of water safety decreased the likelihood of protesting (possibly due to fear), but did not translate into positive bids. Gender played no role in either determining protests or positive bidders.

A multinomial regression model showed that few differences exist between protest and true zeros but that several distinct differences distinguish protest bidders from positive bidders. This indicates that if the respondents who gave a protest bid had not protested, they likely would have given a zero response. The procedure of discarding protest bids when calculating the mean or median WTP assumes that the true responses of those giving protest bids would be distributed like those who did not protest. The results suggest that, at least for the present study, discarding protest bids introduces a bias in the characteristics of the sample used to calculate the mean WTP for groundwater quality protection. While the alternative procedure of counting all protest bids as a zero bid (in order to give a conservative estimate of the total value of a policy or good) would have been appropriate in this case, there is no *a priori* reason to assume that this is always the case. An objective test to determine how to treat protest bids will be useful in policy applications of CVM.

The general inability to differentiate respondents who provided true zeros from those giving protests means that tailoring survey design to reduce protests would be difficult. On the other hand, differences *between* protest bidders and positive bidders indicate that question design is still important in CVM studies. Improving the design of the protest identifier question warrants further study.

Table 1. Bid Classification

Type	response Q22	DOE	IOE	Total	Percent
Positive		129	151	280	45.4
Real Zero	2	22	17	39	6.3
	4	12	16	28	4.5
Total Real Zeros		34	33	67	10.9
Total of Positive Bids and Real Zeros		163	184	347	56.2
Protest	1	21	22	43	6.9
	3	19	32	51	8.2
	5	23	39	62	10.0
	gov/tax	15	21	36	5.8
	other	2	4	6	0.9
Total Protest Zeros		80	118	198	32.1
Missing		41[a]	31	72	11.7
TOTAL		284	333	617	100

[a] 12 of the missing DOE bids were illogical and were coded as missing

Table 2. Logit Regression of Protests versus Legitimate Bidders.

y = PROTEST, Frequency of y = 1: 33.39%

Variable	ML estimate of b(.) (t-value)
MOVE	-0.0046355 (-1.4627)
SAFETY	-0.0069470 (-1.7729)
CONTAMINATION INFO	0.1681595 (0.8200)
SMART	0.0911351 (1.6229)
HIGHER ED	-0.4859841 (-2.0116)
DIFFERENCE	-0.0257719 (-6.9089)
AGE*	0.0095922 (1.9488)
INCOME	-0.0019595 (-0.5548)
MALE	-0.0771723 (-0.3654)
INTERCEPT	-0.9070381 (-1.5034)
Log likelihood:	-353.425
Sample size (n):	617

Table 3. Logit regression of Positive Bidders versus Protests and Zeros

y = POSITIVE, Frequency of y = 1: 47.16%

Variable	ML estimate of b(.) (t-value)
MOVE	0.0055881 (1.7532)
CONTAMINATION INFO	-0.2840106 (-1.3690)
SMART	0.2489274 (3.9791)
DIFFERENCE	0.0384994 (8.4483)
INCOME	0.0146875 (4.0710)
MALE	0.3308085 (1.5113)
HIGHER ED	0.2933686 (1.2644)
AGE	0.0034198 (0.6587)
INTERCEPT	-2.2270412 (-3.6117)
Log likelihood:	-330.657
Sample size (n):	617

Table 4. Multinomial Regression Results

```
y = multi, X(1) = MOVE, X(2) = SAFETY, X(3) = CONINFO, X(4) =
SMART, X(5) = HIEDU, X(6) = DIFFERENCE, X(7) = AGE*,
X(8) = INCOME, X(9) = MALE, X(10) = 1
Model:P(y=0|x) = 1/[1+exp(b(1)'x)+ ..+ exp(b(m)'x)]
      P(y=j|x) = exp(b(j)'x)P(y=0|x), j=1,..,m, where m = 2
```

Variable	ML estimate of b(.)	(t-value)
True zeros relative to protest zeros		
x(1)=MOVE	0.0020933	(0.4582)
x(2)=SAFETY	0.0159059	(2.5181)
x(3)=CONTAMINATION INFO	-0.2400342	(-0.8057)
x(4)=SMART	-0.0392618	(-0.5375)
x(5)=HIGHER ED	0.1132831	(0.3124)
x(6)=DIFFERENCE	-0.0050291	(-1.0761)
x(7)=AGE	0.0015947	(0.2381)
x(8)=INCOME	-0.0010776	(-0.2042)
x(9)=MALE	-0.0654083	(-0.2106)
x(10)=INTERCEPT	-1.9665208	(-2.2043)
Positive bids relative to protest zeros		
x(1)=MOVE	0.0058059	(1.7470)
x(2)=SAFETY	0.0015322	(0.3535)
x(3)=CONTAMINATION INFO	-0.3912649	(-1.8062)
x(4)=SMART	0.2358342	(3.6373)
x(5)=HIGHER ED	0.3220401	(1.3264)
x(6)=DIFFERENCE	0.0356536	(7.8074)
x(7)=AGE	0.0040991	(0.7391)
x(8)=INCOME	0.0140136	(3.7117)
x(9)=MALE	0.3113487	(1.3654)
x(10)=INTERCEPT	-1.9407692	(-2.9310)
Log likelihood:	-506.596	
Sample size (n):	617	

Table 5. Multinomial Logit Model: Positive Bids are Base

```
Positive Base(y=0)Model variables: y = zoink, X(1) = MOVE, X(2) =
SAFETY, X(3) = CONINFO, X(4) = SMART, X(5) = HIEDU, X(6) = DIFF,
X(7) = AGE, X(8) = INC, X(9) = MALE, X(10) = 1
Model:P(y=0|x) = 1/[1+exp(b(1)'x)+ ..+ exp(b(m)'x)]
P(y=j|x) = exp(b(j)'x)P(y=0|x), j=1,..,m, where m = 3
```

Variable	ML estimate of b(.) (t-value)
New zeros relative to positive bids	
x(1)=MOVE	-0.0023016 (-0.5316)
x(2)=SAFETY	0.0079488 (1.2515)
x(3)=CONTAMINATION INFO	-0.4473117 (-1.4928)
x(4)=SMART	-0.3158719 (-3.7812)
x(5)=HIGHER ED	-0.2394156 (-0.7202)
x(6)=DIFFERENCE	-0.0300176 (-5.2106)
x(7)=AGE	0.0011123 (0.1226)
x(8)=INCOME	-0.0177497 (-2.6230)
x(9)=MALE	-0.2836159 (-0.9385)
x(10)=INTERCEPT	1.3379304 (1.6093)
Old zeros relative to positive bids	
x(1)=MOVE	-0.0105103 (-2.0567)
x(2)=SAFETY	-0.0078012 (-1.2195)
x(3)=CONTAMINATION INFO	0.1394119 (0.4548)
x(4)=SMART	-0.2177696 (-2.3206)
x(5)=HIGHER ED	-0.3394706 (-0.9377)
x(6)=DIFFERENCE	-0.0440234 (-7.2839)
x(7)=AGE	0.0142852 (1.4658)
x(8)=INCOME	-0.0057842 (-0.8376)
x(9)=MALE	-0.4140676 (-1.2301)
x(10)=INTERCEPT	0.8109228 (0.9279)
New Protest zeros relative to positive bids	
x(1)=MOVE	-0.0053218 (-1.1560)
x(2)=SAFETY	-0.0020382 (-0.3280)
x(3)=CONTAMINATION INFO	-1.0057552 (-3.0143)
x(4)=SMART	-0.1037976 (-1.0968)
x(5)=HIGHER ED	-0.6043729 (-1.7766)
x(6)=DIFFERENCE	-0.0355565 (-6.1822)
x(7)=AGE	-0.0043708 (-0.4690)
x(8)=INCOME	-0.0038789 (-0.6090)
x(9)=MALE	0.0041720 (0.0133)

```
x(10)=INTERCEPT                 0.9025786  (1.0393)

Log likelihood:        -551.087
Sample size (n):        509
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References

Delavan, Willard. 1997. Valuing the Benefits of Protecting Groundwater from Nitrate Contamination in Southeastern Pennsylvania. Master of Science Thesis, The Pennsylvania State University, University Park, PA., 154 pp.

Epp, Donald J. and Willard Delavan, 1997, "Measuring the Value of Protecting Groundwater Quality: Results and Methodological Findings." In Tenth Interim Report of W-133 Benefits and Costs Transfer in Natural Resource Planning, pp. 119-135.

Halstead, John M., A.E. Luloff, and Thomas H. Stevens. 1992. "Protest Bidders in Contingent Valuation." *Northeastern Journal of Agricultural and Resource Economics*. October 160-169.

Kennedy, Peter. 1993. A Guide to Econometrics. MIT Press. Cambridge, Massachusetts.

Loomis, John B. 1996. "Measuring the economic benefits of removing dams and restoring the Elwha River: Results of a contingent valuation survey." *Water Resources Research*. 32: 441-447.

Mitchell, Robert Cameron; Richard T. Carson. 1989. Using Surveys to Value Public Goods: The Contingent Valuation Method. Resources for the Future. Washington, D.C.

Shultz, Steven D., and A.E. Luloff. 1990. "The Threat of Nonresponse Bias to Survey Research." *The Journal of the Community Development Society*. Volume 21 No. 2.

Appendix: VARIABLE DEFINITION

Q #	Variable Name	Variable Description
1	YEARS	Continuous variable describing how many year respondent has lived in the study area
3	MOVE	Whether the respondent plans to move or not on a scale of 0 to 100 where 0=definitely not and 100=definitely will move
4	WATERSOURCE	Drinking water source, 1=Private well, 2=Community Well, 3=Public Water System, 4=Bottled Water
	PRIVATE WELLS	Equals 1 if Q4= 1, respondents with private wells
7	SAFETY	Scale of from 1 to 100 of perception of household drinking water safety (100 = definitely safe)
8	YTEST	Water tested within the past five years, yes=1, no=0
9	CONCERN FOR SAFETY- SELF FAMILY WITH FAMILY NOTWITH OTHER PEOPLE FUTURE GENERATIONS	Concern for groups of people relative to drinking water safety (yourself, your family living with you, your family living in the study area but not with you, people other than yourself living in the study area, future generations who might live in the study area)
10	PRIORITY- ROADS PUBLIC SCHOOLS COLLEGES POLLUTION PARKS CRIME HEALTH CARE DRINKING WATER	Priority for government spending on potential public policy concerns (improving roads and bridges, improving public schools, improving state technical colleges and universities, reducing air pollution from factories, providing more public parks and recreational facilities, reducing crime, improving public health care, protecting drinking water quality)
11	CONTAMINATION INFORMED (CONTAMINATION INFO)	Whether or not the respondent has previously received information about nitrate contamination in the study area, yes=1, no=0
12	HEALTH CONNECTION (HLTHINFO)	Whether or not the respondent has previously received information about the connection between nitrates and health, MALE
13-19	QUIZ1	See questionnaire for quiz questions. The section had 3 multiple choice multiple answer questions and three true false questions.
	SMART	At least 4/6 correct responses for questions 13-19=1, else=0
21_1	Q21_1	Open-ended follow-up for dichotomous choice Willingness to pay question

21_2	Q21_2	Informed open-ended question
22	WHYZERO	Reason for bidding zero
23	SAFETY WITH	Respondent's perception of drinking water safety if the program were to be implemented
24	SAFETY WOUT	Respondent's perception of drinking water safety with if the program were to be implemented
	DIFFERENCE	Q23 -Q24
25	MALE	Dummy variable where male=1 and female=0
26	BIRTH	Last two digits of the year in which the respondent was born
	AGE	Age in years of person answering questionnaire = 96-Q26.
	CHILD-PRESENT	Children under the age of 4 in the Household , yes=1, No= 0
30	EDUCATION	Level of education (less than high school, high school, some college, college degree, professional degree, other)

Valuation of Groundwater Quality: Contingent Values, Public Policy Needs, and Damage Functions

Gregory L. Poe*

Abstract: In a departure from past contingent valuation research of groundwater quality, this paper estimates a damage function for nitrate exposures based on actual water test results of individual wells. From the perspective of reliability, it is argued that such a full information approach more closely represents the goal of valuation research in this area to estimate the economic values that people would place on improving water quality if they were actually experiencing contaminated water. The adoption of a damage function approach linking willingness to pay to actual exposures is also more useful to policy makers at the study site because it potentially provides benefit information to a broad range of policy options. Finally, because the damage function is based on objective data that could be obtained from other sources such as local well test programs, such an approach may be desirable from a benefits transfer perspective. Damages, as measured by willingness to pay for protecting individual well supplies within a 10 mg/L NO_3-N health standards are estimated to be a concave function of nitrate exposure levels.

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Valuation of Groundwater Quality:
Contingent Values, Public Policy Needs, and Damage Functions

1. Introduction:

In recent years the need for valuation of groundwater resources has been identified as a critical national research and policy issue [USEPA 1990; National Research Council, 1997]. Corresponding to this need, there has been intensive research effort in the last decade to estimate contingent values for groundwater quality [Edwards, 1988; Schultz and Lindsay, 1990; Powell, 1991; Caudill and Hoehn, 1992; McClelland *et al.*, 1992; Sun *et al.* 1992; Jordan and Elnagheeb, 1993; Poe, 1993; Sparco, 1995; Barrett *et al.* 1996; Delavan, 1996; Randall and deZoysa, 1996; Crutchfield *et al.*, 1997]. Over much of the same period, a renewed interest in assessing the accuracy of benefits transfers has emerged [USEPA, 1993; Loomis, 1992; Downing and Ozuna, 1996; Kirchhoff *et al.*, 1997], with some attention paid specifically to water quality [Vandenberg *et al.*, 1995; Crutchfield, 1995; Crutchfield *et al.*, 1997].

This paper argues that there is an inherent incompatibility between groundwater contingent valuation research as it has developed in the last decade, and groundwater management policy and benefits transfer needs. Past contingent valuation groundwater research has provided important, policy relevant information to decision makers. Yet the objective hypothetical exposure (Suppose your home tap water is contaminated by nitrates to a level that exceeds the EPA s minimum standard by 50% , Crutchfield *et al.*) and the subjective risk (How safe do you feel about your household drinking water supply? , Powell) approaches utilized in past research are not directly amenable to the variety of policy outcomes needed to be considered by water managers in studied sites. Moreover, even though values across groundwater studies and sites have been shown to vary in a systematic manner [Boyle *et al.,* 1994], the value information provided by the original studies precludes transfers to unstudied sites unless fairly restrictive assumptions about identical nature of preferences, perceptions, and exposure levels are made. An alternative to meeting these policy needs would be to reorient groundwater contingent valuation research so that the focus is on actual, objectively obtainable, exposure levels experienced at a study site. Towards this goal, this paper provides the results from a groundwater contingent valuation study that tested individual wells for nitrates, and then solicited WTP values for a groundwater protection program.

The organization of the paper is as follows. Section 2 expands on the arguments introduced in the previous paragraph. The third and fourth sections provide a summary of a contingent valuation study of willingness to pay for a rural well water program that maintains nitrate levels within government standards of 10 mg/1 NO_3-N. The critical difference between this and previous CV research is that the values are directly linked to actual exposures as measured by nitrate test results in the studied wells, allowing the estimation of a damage function consistent with theoretical, management and policy needs. The final section discusses the implications of this research.

2. Limitations of Past Groundwater Quality Contingent Valuation (CV) Research:

Since the publication of the Edwards study, a body of CV research has emerged for valuing improvements in groundwater quality. These studies can be categorized by how the valuation scenarios are structured[1]. One group follows Edwards lead by specifying an objective

[1] The McClelland *et al.* study deviates from other research by focusing on quantity shortfalls associated with shutting down contaminated sources, and is thus not included in this categorization.

hypothetical initial exposure condition and an alternative hypothetical improvement [Jordan and Elnagheeb; Sparco; Crutchfield *et al.*; Delavan]. Other studies have allowed respondents to specify their own subjective probabilities of exceeding health standards in a specific time frame [Sun *et al.*; Poe] or perception of current safety levels [Powell], with the target being the reduction of the probability of exceeding standards to zero or the improvement of water quality to safe levels. Still other research has respondents value broadly defined groundwater protection programs and policies [Schultz and Lindsay; Caudill and Hoehn; Randall and deZoysa; Barrett *et al.*]. While providing useful information about willingness to pay for hypothetical programs, each of these approaches has limitations from a valuation, management, or policy perspective. These limitations are discussed here.

The first issue is how well this entire body of groundwater valuation literature represents willingness to pay if the households' water were indeed contaminated and the respondents were actually faced with decisions about public intervention and averting opportunities. To make such decisions, individuals need an adequate amount of information [Arrow *et al.*, 1993; Fischoff and Furby, 1988]. Information gathering has opportunity costs, and individuals may ration scarce information gathering resources by choosing to ignore information that is not relevant to current choices [Bishop and Welsh, 1992]. Such rationing appears to be the norm for specific environmental risks. For example, in a baseline study of radon, about 25 percent of respondents were unable to answer whether their current household exposure was serious or not serious or some level in between [Smith *et al.*, 1990]. With respect to groundwater quality, two water testing studies indicate that most households are unsure about their nitrate exposure levels relative to health standards, and that about 40 percent of rural residents who rely on their own wells are unable to attach a safety level to their water supplies [Poe *et al.*, 1996]. This evidence strongly implies that reliance on subjective perceptions of exposure and health risks may not provide a reliable reference point for valuing a protection policy. People simply do not have well-formed reference conditions, and thus it is unlikely that values collected under these conditions would reliably predict WTP values for a population actually experiencing contamination. The alternative approach of providing participants with an objective hypothetical exposure levels also has limitations. Both the radon and groundwater literature indicates that individuals do update their risk perceptions, and consequently their WTP for protection, with new information. Importantly, they also place weight on their prior perceptions in assessing new information about risks -- even when these priors are erroneous. Given this evidence of updating, it is not known how a household that believes their water to be safe reacts to being asked to assume that their water violates government health standards (or vice versa). At issue is whether adding a hypothetical reference exposure level is meaningful: Do households actually experiencing contaminated water at a given level react similarly to households that are asked to assume that they are experiencing contamination at the same level? At this point in time this question remains unanswered by the CV literature, representing a plausible but yet unquantified bias.

Beyond the reliability of individual values, there is a need to design research so as to provide critical information to groundwater managers and policy makers. A recent National Research Council panel notes that what is most relevant for decision making regarding groundwater pollution policies or management is knowledge of the how economic values will be affected by a decision affecting levels of contamination. This policy perspective reflects, in part, the theoretical requisites for identifying optimal groundwater pollution policies for groundwater, which rest on the notion of damage functions across nitrate exposure levels [e.g. Conrad and Olson, 1992]. Conceptually, it also reflects the necessary information for evaluating the welfare effects of alternative land use practices on the distributions of pollutants [Boisvert, Schmidt, and Regmi, 1997; Wu and Babcock, 1995; Lichtenberg and Zilberman, 1988]. What mangers need in order to meet these policy and managerial issues is information that would allow them to compare the benefits and costs associated with a range of alternative shifts in exposure distributions. To a large extent, past research has been fairly limited with respect to providing such a range of information. Research into specific policies or specific changes in exposures provide little information beyond those specific changes, and thus, has limited relevance to managers interested in exploring a range of alternative programs. The coarse percentages (0, 25, 50, 75, and 100) utilized in much of the objective hypothetical and subjective groundwater research also do not facilitate such comparisons: for example, given a health standard of 10 mg/l NO_3-N, how is a move from 9 to 7 mg/l or from 15 to 10 mg/l NO_3-N to be evaluated? More generally, how are shifts in entire distributions to be assessed? Clearly, for management purposes a damage function approach linking actual exposures to values would be useful for linking social benefits to the control of pollutants.

A second area of policy need is benefits transfers. Following Boyle and Bergstrom [1992] and Desvousges *et al.* [1992], benefit transfers can be defined in the groundwater context as the transfer of existing benefit estimates from an original study site to a change in exposure at an unstudied policy site. The need for such transfers is motivated by relatively high cost and time considerations of conducting original research at the policy site. One way to minimize costs of transfers would be to limit the covariates used in statistical analyses of willingness to pay functions to those that might be readily obtainable from prior research at the policy site: demographic and socio-economic variables (e.g., age, household composition, and income) used in estimating WTP functions could be limited to those corresponding to census records; distributions of groundwater contaminants might be available from hydrologic research in the area [e.g. Portage County Groundwater Plan, 1987; Baker, 1990]. Obviously, studies in which the original research focuses on a localized site-specific issue or policy option will not be likely candidates for benefits transfers. The objective hypothetical or subjectively defined probability also has limited value from a benefits transfer perspective. Given that past research has not linked these values to actual exposure levels, transferring these values to an unstudied site poses a difficulty without conducting a second survey at the study site to determine the range of distributions of probabilities exceeding standards.

In all, from the perspective of obtaining informed values that reflect the best interests of individual decision makers actually experiencing contamination, the need to provide policy makers with valuation data to explore a range of management decisions and the need to conduct benefits transfers, it is argued here that groundwater valuation studies should be based on actual exposures levels and informed respondents. The remainder of this paper describes the first groundwater CV research to be based on actual exposure and to provide a fully informed damage function amenable to local management decisions and benefit transfers.

3. Conceptual Framework

Groundwater valuation of quality changes can be depicted in a standard option price framework [Boyle *et al.*] in which uncertainty is expressed over health states. With respect to nitrates (N) found in well water, the consumer's choice problem can be characterized by the minimization of the planned expenditure function [Smith, 1986]:

$$\dot{e}(g(h;N),p,\overline{EU}) = \min_{X} p'X \quad \text{subject to } EU = \overline{EU} \tag{1}$$

where: $\dot{e}(\cdot)$ is the planned expenditure function; g(h;N)) is the subjective distribution of health outcomes (h) for a given nitrate exposure levels N; p is the corresponding state-independent vector of prices for all goods (X) including the explicit or implicit prices for substitute water sources, and $\overline{EU}$ is the reference level of expected utility. *Ex ante* willingness to pay (i.e., before the health risk is resolved) for a groundwater protection program that shifts the exposure distribution from N to N' is given by the difference in the planned expenditure function with the project and the planned expenditure function without the project:

$$WTP_{N'|N} = \dot{e}(g(h;N),p,\overline{EU}) - \dot{e}(g(h;N'),p,\overline{EU}) \tag{2}$$

More typically however, groundwater protection projects are defined, as in this research, in terms of truncating the nitrate distribution f(N') at some health standard or threshold (T). For example, most nitrate studies to date [e.g., Sun *et al.*; Crutchfield *et al.*] have formulated the target nitrate level in terms of a zero probability of exceeding standards. In this case, the willingness to pay is given by:

$$WTP_{T|N}=\dot{e}(g(h;N),p,\overline{EU}) - \dot{e}(g(h;\int_0^T f(N')dN),p,\overline{EU}) \tag{3}$$

where f(N') depicts a distribution of exposures given the project. Using this expenditure difference, a damage function relative to the threshold level could be obtained from cross-sectional data with varying initial exposure levels. To isolate effects of moving along a damage function, Equation (2) could be approximated by the differencing of Equation (3) across initial nitrate levels under the assumption that the truncated distribution f(N′) is independent of the initial level of exposure and that health risk perceptions across nitrate levels are independent of reference nitrate levels[2,3].

4. Survey Implementation:

The groundwater survey was conducted for private wells in Portage County, Wisconsin, an area known to have a wide range of nitrate distributions based on previous hydrologic research and water testing programs. Prior water testing indicated that approximately 18 percent of the private wells exceed the government health standards of 10 mg/l NO_3-N designed to protect infants from *methemoglobinemia.*

In order to test individual wells and obtain values based on well test results, a two-stage survey design was created. In the first stage (Stage 1), individual households received the following survey package: a cover letter; a Wisconsin State Laboratory of Hygiene mailable nitrate test kit; instructions for collecting a water sample for nitrates; a question and answer sheet providing further information about the study; a business reply return envelope; and an initial survey about respondent

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[2] Both these assumptions may be questionable. For example, it is likely that an individual whose nitrate well test is 2 mg/l will likely have a different perception of f(N') truncated at T = 10 mg/l NO3-N than an individual whose reference nitrate level is 20 mg/l. Similarly, prospect reference theory [Viscusi, 1989] suggests that individuals will formulate perceptions of health risk based on their exposure level. Nevertheless, given that the magnitude of possible biases is not known, it is argued that the willingness to pay values for a shift in distributions as suggested in the text could be used as a rough approximation to evaluate incremental shifts in reference nitrate levels.

[3] In specifying Equation (3), it is of course recognized that, even with the nitrate test results, the reference conditions may also be characterized by a distribution of exposures, say F(N). Previous research, suggests that nitrate levels in individual wells may fluctuate over time [e.g. Baker]. Adding such a redefinition would not change the essence of the analysis -- it merely suggests that the single test approach adopted in this research lies somewhere along the continuum of uninformed to fully informed. Similarly, it is possible to regard f(N') as a normalized distribution wherein the observed CDF F(N') is adjusted to reflect the mass at the truncation point.

socio-demographic characteristics, prior knowledge of groundwater, and safety perceptions. This survey also contained a subjective/uninformed CV question about a 10 mg/l groundwater protection program. Water samples from the Stage 1 respondents were tested for NO_3-N at the Wisconsin State Laboratory of Hygiene. These results were returned to the Stage 1 respondents in a second survey package, which also contained a nitrate information sheet, a Stage 2 CV questionnaire, and a stamped first class return envelope. The information sheets were based on information readily available at the local extension offices and other State and County agencies, and included background information on sources of nitrates, health effects of nitrates, and a listing of possible averting opportunities available to individuals.

The contents of the survey received design input from other CV practitioners and were evaluated in three individual in-person debriefing sessions. The two stage survey design was pre-tested on 20 Portage County households. Based on these pre-tests and other inputs, only minor wording changes were made in the final questionnaire.

Implementation of the survey followed Dillman s total design method [Dillman, 1978], employing an initial survey package, a thank you/reminder post card to all respondents, and a follow-up survey package to those who had failed to reply to the initial survey package. [4] No financial incentives were provided, but participants were informed that the free nitrate test had a $9.00 value.

A zip-code based sample list was obtained from Americalist, and cross checked with local plat books to isolate residences not connected to public water supplies. The survey was initially sent

[4] For Stage 1, in place of Dillman's suggested registered mail third follow-up, telephone contacts were made with survey recipients whose telephone numbers could be identified. A third mailing was sent to those contacted who indicated on the telephone that they would consider completing a questionnaire.

to 480 addresses in rural areas of Portage County that did not have public water supplies. After accounting for bad addresses and addresses outside of the desired area (n = 47), the adjusted Stage 1 response rate was approximately 77% (n = 332). The conditional response rate for the Stage 2 survey was about 83% (n = 275). Each of these individual response rates exceeded the present CV standard of 70%, and the combined response rate across the two stages was about 64%. Even though the 64% response rate reflects non-participation across both survey stages, this ratio still lies at the upper end of the range of single stage groundwater valuation studies [Jordan and Elnagheeb (35%); Barrett *et al.* (45%); Powell (50%); Randall and deZoysa (51%); Sun *et al.* (51%); Schultz and Lindsay (58%); McClelland *et al.* (60%); and Hoehn (66%); Edwards (78.5%)].[5]

Nitrate test results reflected prior water testing results for Portage County. In this study about 16 percent of the wells exceeded government standards of 10 mg/l, with the highest values being 43 mg/l. This corresponds closely with the 18% figure obtained from previous sampling in the area. About 28 percent fell below the highest natural levels of 2 mg/l. The majority of respondents, about 56 percent, had some evidence of human impact on nitrate levels but did not exceed government standards. Thus, a wide range of exposure levels was available to serve as input for a damage function[6].

The two stage questionnaire complicates discussion of the flow of the survey. The first stage

[5] Heckman type selection tests were conducted across stages. Nitrate test levels, demographic and socio-economic variables were included in a probit analysis across stages. Only the age of the respondent (+) and bottled water users (-) were significant factors in explaining whether a Stage 1 respondent completed a Stage 2 questionnaire. However, inverse Mills ratios derived from this analysis were not a significant explanatory variable in estimating Stage 2 willingness to pay response functions, and are, thus, not included in the econometric analysis below.

[6] About 10 percent of the respondents had levels less than measurable (i.e. < 0.15 mg/l) by the techniques used by the Wisconsin State Laboratory of Hygiene. These were excluded from the econometric analyses because they had a special sticker manually placed in their surveys indicating that it was not possible to improve their water quality.

was constructed as a standard stand-alone CV questionnaire, obtaining information about personal perceptions of groundwater exposure and health risks, eliciting other background information on respondents' environmental concerns, eliciting a yes/no response to a dichotomous choice CV question for a 10 mg/l standard based on pre-existing, subjective/ uninformed values, and then obtaining socio-economic descriptors. The second stage questionnaire focused instead on personal impressions of their individual water test results and, given that information, the relative safety of their water. Individual averting options were discussed and a community-wide program was presented as an alternative to individual protection. Following a reminder that taxpayers, individuals, and farmers already pay for groundwater protection through government programs, higher prices, and lower profits, the following program was proposed:

- *<u>With</u> the groundwater protection program, nitrate levels in all Portage County wells will <u>definitely</u> be kept below the government health standard of 10 mg/l. In some areas this may be difficult, but suppose that it would be possible.*

- *<u>Without</u> such a groundwater protection program, present trends in nitrate levels in Portage County will continue and the number of wells with nitrate levels higher than the government standard will increase in Portage County in the next five years.*

Respondents were subsequently asked to vote in a subjective/informed or fully informed manner on the program with the following dichotomous choice contingent valuation question:

> *Would you vote for the groundwater protection program described above if the total <u>annual</u> cost to your household (in increased taxes, lower profits, higher costs, and higher prices) were $________ each <u>year</u> beginning now and for as long as you live in Portage County (CIRCLE ONE NUMBER)*
>
> *1. No*
> *2. Yes*

Dollar values were individually inscribed and ranged from $1 to $999. The range and distribution of these bid values were based on information obtained from the Stage 1 survey responses.

5. Econometric Methods

Corresponding to the expenditure approach described in Equations (1) to (3), estimation of the WTP function follows the expenditure difference random utility model initially described by Cameron [1988, 1991; McConnell, 1990]. In this framework, the possibility of a 'yes' response to the dichotomous choice bid value 'A' is given as:

$$\pi(yes) = \pi(WTP_{TJN} + \varepsilon \geq A) \tag{4}$$

where the error term is assumed to have a zero mean. WTP_{TJN} is unobserved but indicated by the 1/0, yes/no response to the dichotomous choice question. Assuming a logistic distribution for ε the following relation provides a first step in recovering an estimated WTP_{TJN} function:

$$\pi(yes) = (1 + \exp(-(\alpha + \beta A + \underline{\gamma}\underline{Z})))^{-1} \tag{5}$$

where $\underline{Z}$ is a vector including a function of nitrate levels and demographic characteristics of the respondent, and α,β, and $\underline{\gamma}$ are coefficients to be estimated. Estimated WTP_{TJN} for an individual can be recovered by the following transformation:

$$WTP_{TJN} = \frac{\alpha}{\beta} + \frac{\underline{\gamma}}{\beta}\underline{Z} \tag{6}$$

Derivation of standard errors for the ratios of coefficients follows the standard logistic estimation procedures detailed in Cameron's 1992 article.

In the statistical analyses that follow, Z will be defined to consist of two components. The first component contains of socio-demographic variables of the type that could be linked to census type data for benefits transfers. These covariates, and their expected correlation with WTP, are: the age (-) and gender (?) of the respondent; presence of children less than 4 years of age in the household (+); involvement in farming (-); education level (+); and household income (+). These variables are further defined in Table 1. Expectations of the sign of the estimated coefficients were taken from other CV research on valuing risks.[7]

Importantly, from the perspective of this paper, Z also includes a nitrate exposure variable, for which the derivation of the conditional WTP is the objective of this research. Two approaches to characterizing exposure levels are evaluated. The first corresponds with the subjective/informed probability of exceeding standards approach. Immediately preceding the valuation question, the following question about exposures was posed:

> *Without such a groundwater protection program, do you expect that your own well will have more nitrates than the government standard of 10 mg/l during the next five years? If you are not sure, please give us your best guess. (CIRCLE ONE NUMBER)*

> ***1. Yes, my well already has more nitrates than the 10 mg/l standard and I expect it to remain above the standard.***
> ***2. Yes, definitely (100 percent chance)***
> ***3. Probably (75 percent chance)***
> ***4. Maybe (50 percent chance)***
> ***5. Probably not (25 percent chance)***
> ***6. No, definitely not (0 percent chance)***

Responses to this question were recoded according to their probability of exceeding standards to

[7] See Poe and Bishop (1997) for a more detailed discussion of these variables.

form the covariate Pr(NO3-N > 10 mg/l) with a range between 0 and 1 in 0.25 increments. The expectation is that the coefficient on this variable would be positive, reflecting the well established result that people with a higher perceived likelihood of exposure will have a greater WTP for protection. Given this formulation, there is no direct link to exposure levels. Although, as discussed below, such a relationship might be obtained by linking expectations to exposures in a secondary analysis.

The second approach instead focuses on establishing a direct damage function relationship between WTP responses for the 10 mg/l protection program and nitrate levels. Little prior empirical evidence exists about the shape of this function. All else equal, we would expect that people with low reference exposures would have low WTP for a protection project, while households with high exposures would have a relatively high WTP for such a project. However, when linking WTP directly to exposures, concern must be given to the convexity of the damages between these two extremes. On one hand, the standard value of life literature would suggest a convex damage function [Jones-Lee, 1974]. However, when substitutes or defensive expenditures such as bottled water are included as decision options, the damage function may become non-convex [Burrows, 1995]. In all, convexity of the damage function is an empirical question [Shogren and Crocker, 1991; Quiggen 1992]. Ignoring for the moment all other elements of Z, convexity is investigated by assuming the following reformulation of Equation 6:

$$WTP_{TJN} = \frac{\alpha}{\beta} + \frac{\gamma_N}{\beta} N^{\tau} \tag{7}$$

In this specification, $\tau>1$ implies a convex damage function and $\tau<1$ corresponds to concavity. In the analyses that follow, an optimal τ is determined by a grid search with the objective of minimizing the likelihood function. Once determined, τ is fixed and the remaining coefficients are estimated using standard logistic maximum likelihood techniques.

6. Results

The results of the estimation process are summarized in Table 2. The first column of the table defines the coefficient or the variables to be estimated. The second column provides the mean values and standard deviations for relevant variables. The third through fifth column reports coefficients and estimated summary statistics for maximum likelihood estimates corresponding to Equations (4) through (6). Different columns in this set correspond to different specifications of $\underline{Z}$. In the first specification, the Pr(NO3-N >10 mg/l) is the only variable in $\underline{Z}$. The second specification expands the definition to include all the socio-economic variables except income. The third specification includes income as an element of $\underline{Z}$, at the cost of losing about 10 percent of the observations. The final three columns of Table 2 report the model demand function defined by Equation 7 for the same sequence of covariates.

Within each specification of the nitrate variable, the three formulations of $\underline{Z}$ exhibit similar trends. Coefficients on the nitrate variables are highly significant, with appropriate signs in all specifications. In the estimates excluding INCOME, the coefficients on OWNAGE and DCGRAD are negative and positive respectively, as expected. The other coefficients are not significant. When INCOME is included, all the coefficients for the remaining non-nitrate covariates become insignificant. This suggests that estimation of a WTP function will be dominated by income and the level of exposure. Should this result be supported by future research, benefits transfers might be accomplished by relatively simple models of income and exposure.

Although both specifications are significant at the 1% level, a comparison of the informed subjective probability models with the corresponding nitrate exposure model indicates that the former provides a better statistical estimate of WTP: the variance of the WTP estimate (given by κ) is smaller, the χ^2 goodness of fit statistic is higher, and the percent of responses correctly predicted is higher. Thus, if the sole objective is goodness of fit, then the subjective/informed approach based on the likelihood of exceeding standards would dominate.

However, as discussed in Section 2, such an approach is limited by its indirect linkage to nitrate levels. From the perspective of local management policies and the potential for benefits transfers, it is policy useful to have WTP estimated as a function of nitrate levels. Such a direct estimate is provided in the last three columns of Table 2. In this analysis, $\frac{\gamma_N}{\beta}N^K$ indicates an increasing concave function of nitrates. That is, WTP_{TIN} rises with N but in a decreasing manner. Given the grid search approach adopted here, direct statistical tests of concavity cannot be performed. However, support for this conclusion is found by bootstrapping the data set and identifying an optimal τ for each bootstrap sample. Using this approach, 87 of 100 bootstrap estimations provided τ values of less than 1.

Figure 1 provides a graphical depiction of this damage function based on the simple model Z model in Column 6 of Table 2. As depicted, the direct nitrate exposure model provides a concave function that levels off at higher reference exposure levels. Such a result is consistent with opportunities for substitutes (part of the information packet provided with the Stage 2 survey). Taking averages of expectations about the probability of exceeding standards across ranges of nitrate levels provides point estimates at various levels of a derived damage function. In contrast to the concave damage function, the estimated damages rise relatively slowly across low levels of NO_3-N contamination, jump sharply as reference exposures cross 10 mg/l, and then level off as the expectations of exceeding approach 100 percent. The resulting damage mapping suggests an 'S' shaped function of damages, wherein a convex function corresponding to standard value of life hypothesis occurs across lower values, but the WTP values are eventually truncated from above.

7. Discussion:

This paper suggests that CV research on groundwater quality and other environmental risks adopt a paradigm that WTP values should be based on actual exposure levels. Arguments underlying such a proposal center on the reliability of individual WTP responses as well as the need to provide land use and groundwater managers and policy makers with valuation data that can be linked to a range of decisions. Such an approach would also provide more flexible input for benefits transfers.

Towards this objective, this paper provides the results from the first CV survey of groundwater nitrate contamination to be based on actual exposure levels experienced by respondents. Willingness to pay for a program to protect groundwater at a 10 mg/l NO3-N standard was obtained from respondents who had been informed of their households' nitrate test results. Adopting an expenditure difference approach, a damage function was estimated linking WTP to actual exposure levels. In analyzing the dichotomous choice response, a relatively simple functional form for nitrates was estimated within a logistic framework, resulting in convex damages. An indirect approach, obtained by first estimating WTP as a function of subjective probabilities of exceeding standards and then linking these probabilities to exposure levels, suggests a damage function with convexities and concavities. Nevertheless, in contrast to standard presentations of damages, both approaches suggest that WTP eventually levels off. Such a result is consistent with opportunities for substitution. Examination of more sophisticated functional forms remains a critical area of future research.

In arguing that a fully informed approach should serve as the paradigm for future research, it is recognized that testing water quality may be expensive, perhaps prohibitively so in some situations. Nevertheless, it is incumbent upon researchers, policy makers, and funding agencies to recognize that values based on partial information will provide limited, and perhaps biased, information to decision makers. The benefits of obtaining values from a fully informed sample are likely to be more than marginal, and thus merit consideration in future policy relevant research.

References:

Arrow, K., R. Solow, E. Leamer, P. Portney, R. Rander and H. Schuman, 1993. Report of the NOAA Panel on Contingent Valuation , Natural Resource Damage Assessments Under the Oil Pollution Act of 1990, NOAA, U.S. Department of Commerce, Federal Register, 58(10): January 15, 4602-14.

Baker, D. B., 1990. Groundwater Assessment through Cooperative Private Well Testing: An Ohio Example . Journal of Soil and Water Conservation, 45(2): 230-35.

Barrett, C., T. H. Stevens and C. E. Willis, 1996. Comparison of CV and Conjoint Analysis in Groundwater Valuation , W-133 Benefits and Costs Transfer in Natural Resource Planning, Ninth Interim Report, Western Regional Research Publication, July, pp. 79-129.

Bishop R. C. and M. P. Welsh, 1992. Existence Values in Benefit-Cost Analysis and Damage Assessment , Land Economics, 68(4):405-417.

Boisvert, R. N., A. Regmi and T. M. Schmit, 1997. Policy Implications of Ranking Distributions of Nitrate Runoff and Leaching from Corn Production by Region and Soil Productivity , Journal of Production Agriculture, forthcoming.

Boyle, K. J. and J. C. Bergstrom, 1992. Benefits Transfer Studies: Myths Pragmatism and Idealism . Water Resources Research, 28(3): 657-63.

Boyle, K. J., G. L. Poe, and J. C. Bergstrom, 1994. What Do We Know About Groundwater Values? Preliminary Implications from a Meta Analysis of Contingent Valuation Studies , *American Journal of Agricultural Economics*, 75(5):1055-1061.

Burrows, P., 1995. Chapter 12: Non-Convexities and the Theory of External Costs , *The Handbook of Environmental Economics*, D. W. Bromley, ed. Basil Blackwell Ltd., Oxford.

Cameron, T. A., 1988. A New Paradigm for Valuing Non-Market Goods Using Referendum Data: Maximum Likelihood Estimation by Censored Logistic Regression , Journal of Environmental Economics and Management, 15, 355-79.

Cameron, T. A., 1992. Interval Estimates of Non-Market Resource Values from Referendum Contingent Valuation Surveys , Land Economics, 67(4): 413-21.

Caudill, J. D. and J. P. Hoehn, 1992. The Economic Valuation of Groundwater Pollution Policies: The Role of Subjective Risk Perceptions , Staff Paper No. 92-11, Department of Agricultural Economics, Michigan State University.

Conrad, J. M. and L. J. Olson, 1992. The Economics of a Stock Pollutant , Environmental and Resource Economics, 2: 245-58.

Crutchfield, S. R., 1995. Estimating the Value of Ground Water Protection: An Application of Benefits Transfer , Selected paper, American Agricultural Economics Association Meeting, San Diego, Aug. 7-10.

Crutchfield, S. R., J. C. Cooper and D. Hellerstein, 1997. Benefits of Safer Drinking Water: The Value of Nitrate Reduction , Agricultural Economic Report Number 752, USDA-ERS, Washington, D.C.

Delavan, W. A., 1996. Valuing the Benefits of Protecting Groundwater from Nitrate Contamination in Southeastern Pennsylvania, Master's thesis, The Pennsylvania State University.

Desvousges, W. H., M. C. Naughton, and G. R. Parsons, 1992. Benefits Transfer: Conceptual Problems in Estimating Water Quality Benefits Using Existing Studies, Water Resources Research, 28(3): 675-83.

Dillman, D. A., 1978. Mail and Telephone Surveys--The Total Design Method, New York: Wiley & Sons.

Downing, M. and T. Ozuna, 1996. Testing the Reliability of the Benefit Function Transfer Approach , Journal of Environmental Economics and Mangement, 30(3:May): 316-322.

Edwards, A. F., 1988. Option Prices for Groundwater Protection , Journal of Environmental Economics and Management, 15: 465-87.

Fischoff, B. and L. Furby, 1988. Measuring Values: A Conceptual Framework for Measuring Values with Special Reference to Contingent Valautaion of Visibility , Journal of Risk and Uncertainty, 1:147-84.

Jones-Lee, M. , 1974. The Value of Changes in the Probability of Death or Injury , Journal of Political Economy, 82:835-49.

Jordan, J. L. and A. H. Elnagheeb, 1993. Willingness to Pay for Improvements in Drinking Water Quality , Water Resources Research, 29(2): 237-45.

Kirchhoff, S., B. G. Colby, and J. T. LaFrance, 1997. Evaluating the Performance of Benefit Transfer: An Empirical Inquiry , Journal of Environmental Economics and Management, 33(1: May):59-74.

Lichtenberg, E. and D. Zilberman, 1988. Efficient Regulation of Environmental Health Risks, Quarterly Journal of Economics, 103: 167-78.

Loomis, J. B., 1992. The Evaluation of a More Rigorous Approach to Benefit Transfer: Benefit Function Transfer , Water Resources Research, 28(3): 675-683.

McClelland, G. H., W. D. Schulze, J. K. Lazo, D. M. Waldman, J. K. Doyle, S. R. Elliott and J. R. Irwin, 1992. Methods for Measuring Non-Use Values: A Contingent Valuation Study of Groundwater Cleanup , Office of Policy, Planning and Evaluation, U.S. Environmental Protection Agency, October.

McConnell, K. E., Models for Referendum Data: The Structure of Discrete Choice Models for Contingent Valuation , Journal of Environmental Economics and Management, 18, 19-35.

Mitchell, R. C. and R. T. Carson, 1989. Using Surveys to Value Public Goods: The Contingent Valuation Method, Washington, D.C.: Resources for the Future.

National Research Council, 1997. Valuing Ground Water: Economic Concepts and Approaches, National Academy Press, Washington, D. C.

Poe, G. L., 1993. Information, Risk Perceptions, and Contingent Values: The Case of Nitrates in Groundwater , Ph.D. thesis, Department of Agricultural Economics, University of Wisconsin-Madison.

Poe, G. L. and R. C. Bishop, 1997. Valuing the Incremental Benefits of Groundwater Protection When Exposure Levels are Known , Unpublished Manuscript, Department of Agricultural, Resource, and Managerial Economics, Cornell University.

Poe, G., H. van Es, T. VandenBerg and R. Bishop, 1996, Do Participants in Well Water Testing Programs Update Their Exposure and Health Risk Perceptions? , Working Paper 96-11, Department of Agricultural, Resource, and Managerial Economics, Cornell University.

Portage County Groundwater Plan, 1987, Vol. 1., Portage County, Wisconsin.

Powell, J. R., 1991. The Value of Groundwater Protection: Measurement of Willingness To Pay Information and Its Utilization by Local Government Decision Makers , Master's thesis, Cornell University.

Quiggen, J. 1992. Risk, Self Protection and Ex Ante Economic Value Some Positive Results , Journal of Environmental Economics and Management, 23(1): 40-53.

Randall, A. and D. deZoysa, 1996. Groundwater, Surface Water, and Wetlands Valuation for Benefits Transfers: A Progress Report , W-133 Benefits and Costs Transfer in Natural Resource Planning, Ninth Interim Report, Western Regional Research Publication, July, pp. 79-129.

Schultz, S. D. and B. E. Lindsay, 1990. The Willingness to Pay for Groundwater Protection , Water Resources Research, 26(9): 1869-75.

Shogren, J. F. , and T. C. Crocker, 1991. Risk, Self-Protection, and Ex-Ante Economic Value , Journal of Environmental Economics and Management, 20(1):1-15.

Smith, V. K., 1986. Nonuse Values in Benefit Cost Analysis, Southern Economic Journal, 59: 19-26.

Smith, V. K., W. H. Desvousges, F. R. Johnson, and A. Fisher, 1990. Can Public Information Programs Affect Risk Perceptions , Journal of Policy Analysis and Management, 9(1): 41-59.

Sparco, J., 1995. Marginal Valuation of Health-Related Attributes of Ground Water Using Conjoint Analysis . Selected paper presented at the Northeastern Agricultural and Resource Economics Meetings, Burlington, June.

Sun, H., J. C. Bergstrom, and J. R. Dorfman, 1992. Estimating the Benefits of Ground Water Contamination Protection , Southern Journal of Agricultural Economics, 19(4): 63-71.

USEPA, 1990. Reducing Risk: Setting Priorities and Strategies for Environmnetal Protection , Science Advisory Board, Relative Risk Reductions Strategies Committee, USEPA, Washington D.C.

USEPA, 1993. Benefits Transfer: Procedures, Problems, and Research Needs, Proceedings of the June 1992 Association of Environmental and Resource Economists Workshop, Snowbird, Utah. EPA 230-R-93-018, Washington, D. C.

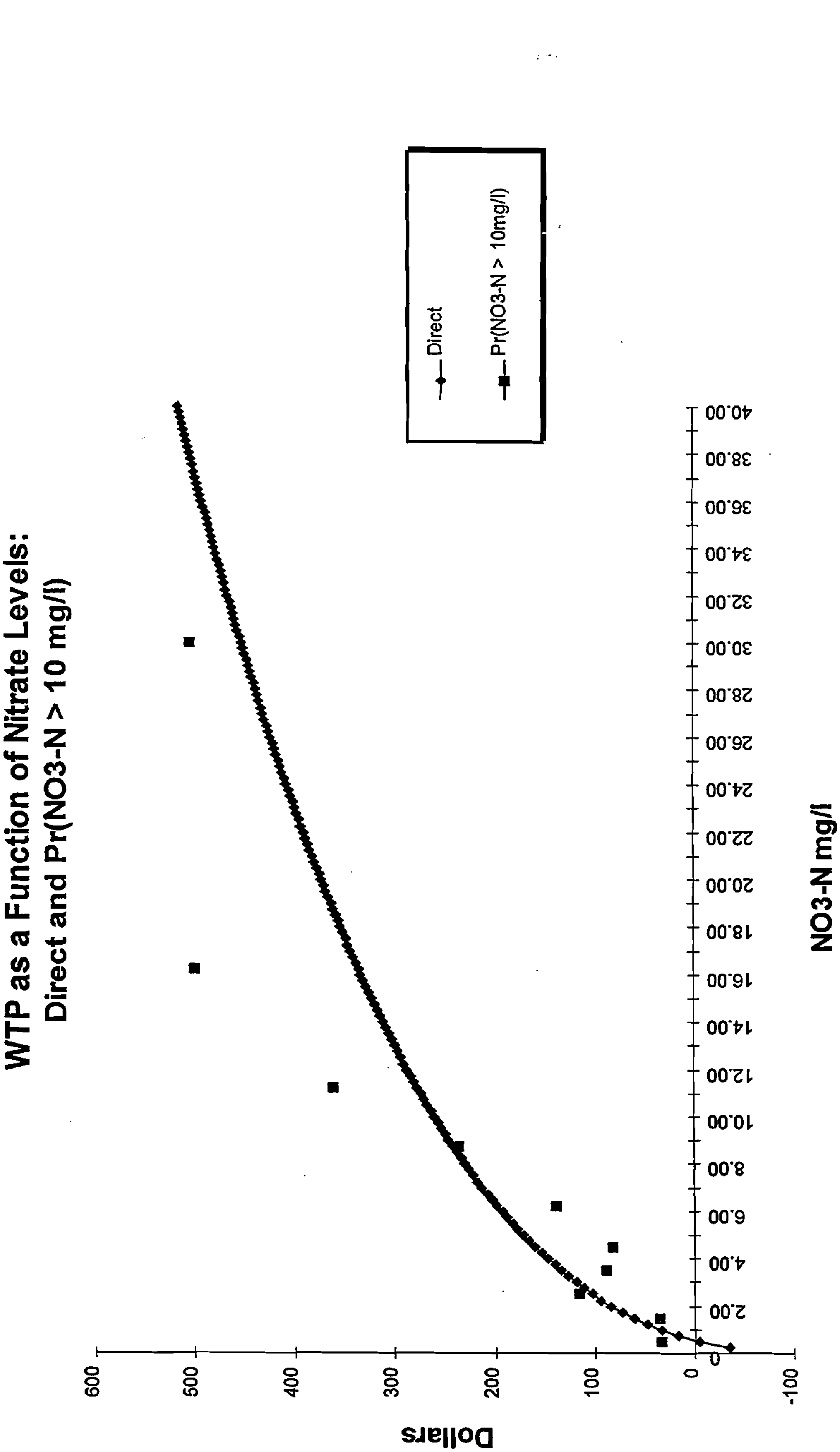
Vandenberg, T. P., G. L. Poe and J. R. Powell, 1995. Assessing the Accuracy of Benefits Transfers: Evidence from a Multi-Site Contingent Valuation Study of Groundwater Quality , selected paper American Agricultural Economics Association meetings, Indianapolis. Working Paper 95-01, Department of Agricultural, Resource, and Managerial Economics, Cornell University. Selected

Viscusi, W. K., 1989. Prospective Reference Theory: Towards an Explanation of the Paradoxes , Journal of Risk and Uncertainty, 2:235-64.

Wu, J. and B. Babcock, 1995. Optimal Design of a Voluntary Green Payment Program Under Asymmetric Information, Journal of Agricultural and Resource Economics, 20: 316-27.

Table 1: Description of the Covariates for The Econometric Analysis

Variable	Description	Sign Expectation
OWNAGE	Categorical Variable for Years of Age: 1= less than 18; 2 = 18 to 44; 3=45 to 64; 4 = 65 or older.	-
DGENDER	Binary variable for gender of respondent: 0= male; 1= female.	?
DAGE<4	Binary variable for young children < 4 years of age in household: 0=no; 1=yes.	+
DFARM	Binary variable for involvement in farming:: 0=no; 1= yes.	-
DCGRAD	Binary variable for college grad: 0=no; 1=yes.	+
INCOME	Categorical variable for total household income before taxes: 1= < $10,000; 2=$10,000 to $19,999; 3=$20,000 to 29,999...10=$90,000 to 100,000; 11= >$100,000	+
P(NO3-N> 10 mg/l)	Probabilistic categorical variable: 0, 0.25, 0.50, 0.75, and 1.00 probability of exceeding standards.	+
N	Nitrate Level (NO3-N) in mg/l, continuous from 0.15 mg/l.	+



WTP as a Function of Nitrate Levels:
Direct and Pr(NO3-N > 10 mg/l)
Dollars
600
500
400
300
200
100
0
-100
0
2.00
4.00
6.00
8.00
10.00
12.00
14.00
16.00
18.00
20.00
22.00
24.00
26.00
28.00
30.00
32.00
34.00
36.00
38.00
40.00
NO3-N mg/l
Direct
Pr(NO3-N > 10mg/l)

Table 2: **Subjective Probability of Exceeding Standards and Nitrate Exposure Models**[a]

Variable	Mean [b] (N= 185)	Subjective Probability Model[c]			Nitrate Exposure Model[c]		
1	2	3	4	5	6	7	8
Constant	1 [0.00]	-102.30 (91.02)	166.51 (203.26)	-284.29 (232.56)	-146.63 (168.99)	237.37 (246.55)	-206.94 (271.27)
OWNAGE	2.67 [2.76]		-108.81 (63.87)*	-21.67 (59.86)		-147.58 (68.52)*	-46.94 (61.16)
DGENDER	0.38 [0.48]		-45.77 (88.95)	46.46 (83.53)		-14.33 (95.28)	43.15 (87.45)
DAGE<4	0.18 [0.39]		-35.22 (114.46)	-8.93 (106.04)		-44.46 (126.58)	8.57 (112.25)
DFARM	0.20 [0.40]		103.67 (104.24)	117.22 (98.69)		61.31 (113.13)	117.14 (104.43)
DCGRAD	0.25 [0.43]		289.54 (106.60)***	99.08 (104.18)		293.70 (115.24)**	109.18 (107.02)
INCOME	4.07 [2.07]			58.49 (22.90)**			55.77 (24.20)**
Prob (>10)	0.45 [0.35]	618.07 (166.96)**	729.38 (195.75)***	528.51 (138.62)***			
τ					0.352	0.353	0.346
N^{τ}					180.75 (89.34)**	146.69 (85.18)*	131.11 (77.92)*
κ^{e}	265.71[d] [276.76]	282.79 (56.36)***	253.91 (48.17)***	213.85 (39.73)***	309.39 (65.03)***	287.24 (58.16)***	234.91 (45.64)***
Obs.	185	210	210	185	210	210	185
χ^2		53.08***	67.85***	72.96***	36.12***	49.42***	55.67
Perc. Pred.		77	73	78	69	73	72

a. ***, **, * denote 1, 5 and 10 percent significance levels, respectively.
b. Numbers in [ ] are standard deviations.
c. Numbers in ( ) are asymptotic standard errors.
d. Mean and standard deviation for the dichotomous choice bid value.
e. $\kappa = 1/\beta$ following Cameron.

Benefits Transfer in RPA: Current Projects

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BENEFITS TRANSFER IN RPA: CURRENT PROJECTS

The objectives of this paper are 1) to provide a brief background on two projects being funded by the Resources Planning Act (RPA) staff, 2) to restate several issues in benefits transfer of nonmarket recreation value estimates for public lands, and 3) to provide some summary statistics from the values database update project. The literature on past benefits transfer efforts provides important insights for future efforts, including the issues addressed, problems identified, and suggestions made. Also, the literature, as documentation of these past efforts, provides a framework or foundation upon which future endeavors can be built.

Benefits Transfer

Benefits transfer is a process by which benefit estimates, or benefit functions, developed for a "study site" are transferred, or adjusted, for application to a "policy site". A study site is a site for which data exist, typically through primary collection techniques. A policy site is a site for which little or no data exist. Traditionally, point estimates have been transferred. However, more recently the transfer of whole benefit or demand functions is favored. Benefit function transfers have the distinct advantage over point estimate transfers in that they are more capable of accounting for significant differences in the site, population demographics, and social institutions between the study and policy sites. These differences between study and policy sites can result in large biases in point estimates.

Performing benefits transfers is justified based on constraints to doing primary research – time and resources (staff and budget). Primary research is expensive and time consuming. In many policy settings, benefit information is needed quickly, negating the possibility of doing primary research. These transfers can be on three levels – the site-to-site transfer, the sites-to-region transfer, and the region-to-site transfer.

Several requirements for performing effective and efficient benefits transfers have been identified in the literature. Desvousges, Naughton, and Parsons (1992) provide five criteria for accomplishing benefits transfers. For RPA purposes, two more criteria need to be stated.

- Studies transferred must be based on adequate data, sound economic method, and correct empirical technique (Freeman 1984).
- The change in the resource quality at the study site and the expected change at the policy site should be similar.
- The study contains regression results of the relationship between benefits and socioeconomic characteristics for the study site.
- The study contains regression results of the relationship between benefits and site characteristics for the study site.
- The markets for the study site and the policy site are similar, unless there is enough usable information provided on own and substitute prices.
- The recreation activity should be similar between the study and policy sites.
- An adequate number of individual studies on a recreation activity for similar sites must be available.

Current Projects

This paper is directly a result of an RPA staff project. The RPA staff is responsible for developing resource values for Forest Service planning. The current project is two-fold. First, the RPA Values Update Project is being undertaken to update the recreation values developed for the 1990 RPA Program. This project is being conducted by John Loomis of Colorado State University. This update is needed to take advantage of the valuation research over the past 10 years (the values are current up to about 1988 research). Currently, RPA provides average values by recreation activity. It would be useful to be able to provide more detailed information about the studies "behind" these average values. In addition to updating the RPA values, the first project will code detailed information on added studies.

And second, the RPA Benefits Transfer Project is being undertaken to explore the feasibility of developing benefits transfer functions for valuation applications. The objective of this second project is to develop benefits transfer functions for RPA recreation activities using meta-analytic methods. These functions will account for systematic effects of methodological differences, user population differences, and differences in the recreation sites on the values reported in the original studies using indicators of these differences that are readily accessible to and estimable by resource managers. This will provide more flexibility and precision in the benefits transfer process that is lacking in the average value approach.

RPA Benefits Transfer Functions

A prototype benefits transfer function can be developed as follows. A single activity is chosen to test the feasibility for and specification of a benefits transfer function. The primary issue is coding measures of differences in valuation methodology, user population characteristics, and site characteristics, enabling the identification of and adjustment for systematic effects from these differences across the original studies. Requirements for selecting a recreation activity include numerous past studies on the value of this activity across a wide geographic range. The generic function will be of the following form:

$$Val|x = f(SE, ST, M)$$

where *Val* is the value reported of the original study per measurement unit *x*, *M* is a matrix of variables accounting for the methodological differences across the original studies, *SE* is a matrix of variables accounting for the socioeconomic (user population) characteristic differences across the original studies, and *ST* is a matrix of variables accounting for the site characteristic differences across the original studies. These matrices of variables will be used to explain and adjust the RPA average value for the recreation activity due to observed differences in the original studies' estimates. The variables selected for inclusion in the equation will be quantitative or qualitative measures that are easily accessible to or estimable by resource managers. The final specification and functional form of this equation will be tested and developed over the life of this project. Several studies will have multiple values when recreation activity types are aggregated (e.g., hunting), necessitating an unbalanced panel data estimation approach.

Methodological Differences

Differences in methodological approaches and model specification in original studies may have systematic effects on the value estimated. The values database will code for these methodological differences including variables such as omission of travel time, individual versus zonal travel cost approaches, stated versus revealed preference expression, in-state versus out-of-state users, inflation, contingent valuation elicitation approaches.

User Population Differences

Differences in user population characteristics surveyed in the original studies may have systematic effects on the value estimated. The ability to correct or adjust for user population differences in a benefits transfer function would enable a more precise and accurate transfer of these values to other sites. The socioeconomic variables to include will be chosen based on theory and empirical results. Potentially important variables could be income, age, sex, household size, education, and other taste and preference indicators. The data for these variables will be collected:

- from original studies when reported,
- using average values from census data as a proxy for the user population at the time of the original study, and/or
- development and use of other proxies.

Site Differences

Differences in site characteristics investigated in the original studies may have systematic effects on the value estimated. The ability to correct or adjust for site characteristic differences in a benefits transfer function would enable a more precise and accurate transfer of these values to other sites. The site variables to include will be chosen based on theory and empirical results. Potentially important variables could be measures of crowding, location, site quality, range of available activities, range and location of substitute sites, etc. The data for these variables will be collected:

- from information included in original studies, or
- through ex post expert assessment by researchers or individuals familiar with the study site, and/or
- development and use of other proxies.

Problems Identified in the Literature

Several limitations to developing effective and efficient benefits transfers have been identified in the literature. Although the majority of these limitations are based on efforts to apply site-specific and user-specific benefit estimates from a study site to a policy site, identifying these limitations is important for future benefits transfer endeavors. Brookshire and Neill (1992) summarized a majority of the problems encountered in other research efforts. The following list of problems is collected from the studies listed in the bibliography at the end of this report.

- *Garbage-in, garbage-out.* The quality of the original study greatly affects the quality of the benefits transfer process.
- *A lack of adequate data points.* Some recreation activities may have a limited number of studies investigating their value.
- *Time can influence values.* The existing studies occurred at different points in time. The relevant differences between then and now may not be identifiable nor measurable based on the available data.
- *Limited data collection in past studies.* Not all of the original studies collected data for the purpose of demand estimation.
- *Unique conditions.* Some of the existing studies may be based on valuing recreation activities at unique sites and under unique situations.
- *Unidentified markets.* The relevant market sizes between the study site and the policy site may not be identifiable nor comparable.
- *Sites may not be comparable.* Characteristics of the study site and policy site may be substantially different, leading to quite distinct values. This can include differences in quality changes, site quality, and site location.

- *User populations may not be comparable.* Characteristics of user populations for the study site may be significantly different than the policy site, including characteristics such as socioeconomic levels, distances from the site, combinations of activities available on the site, and tastes and preferences.
- *Applied research methods.* Different research methods may have been used across study sites for a specific recreation activity.
- *Statistical estimation methods.* Different statistical methods for estimating models can lead to large differences in values estimated. This also includes issues such as the overall impact of model mis-specification and choice of functional form on value estimates.
- *Substitute prices.* There is often a lack of data collection and or reporting on the availability of substitute sites and substitute site prices.
- *Types of values.* There are different types of values (such as use vs. passive-use values) that may be measured.
- *Guidelines.* There are no clear guidelines set for judging the adequacy and scientific soundness of existing studies or of benefits transfer exercises.
- *Purpose of original study.* The existing studies were not designed for benefits transfer purposes.

The above problem areas can lead to bias or error in the benefits transfer process. One objective in the benefits transfer process is to minimize mean square error between the true value and the estimated (transferred) value. However, this assumes the original, or study site benefit estimate is true. Greater awareness of the potential sources of error between true and estimated benefits will assist the benefits transfer practitioner to systematically remove or reduce these biases. The following sources of error are identified primarily by Brookshire and Neill (1992) and by McConnell (1992).

Brookshire and Neill (1992) and McConnell (1992) identified the following sources of error in demand estimation:

- choosing the wrong functional form;
- incompletely or inappropriately specifying the demand function;
- not measuring the variables correctly;
- measuring the dependent variable with gross error; and or
- misspecifying the random process that generates the data.

Brookshire and Neill (1992) and McConnell (1992) identified the following sources of error in benefits estimation:

- incorrect handling of the random component of demand functions;
- aggregation errors;
- incorrect estimation of the affected population;
- incorrect choice frameworks;
- counting trips for multiple purposes; and or
- error in the transfer process.

The next section presents several recommendations made in the literature, primarily in response to the problems identified in this section. These recommendations are targeted at either specific problems or groups of problems, or are generally intended to improve the benefits transfer process as a whole.

Suggestions for Future Benefits Transfer Efforts

Several of the researchers provided recommendations for improving the benefits transfer process based on their experiences. This section lists these recommendations.

- A better (more consistent) reporting and analysis of original study results is needed, including the reporting of unit of output, definition of trip/visit, length of stay and party size, travel time (and value of travel time), and spatial limits to sampling and statistical estimation (Sorg and Loomis 1984; Brookshire and Neill 1992; Boyle and Bergstrom 1992).
- Model specification of original studies needs to be increased to account for important taste and preference variables and site quality variables that are necessary for critical benefits transfers (Loomis, Provencher, and Brown 1990).
- A national clearinghouse of nonmarket valuation studies and a national values database needs to be organized, using some set of standards for valuation and reporting of results and variables, thus standardizing data collected and making it more accessible (Boyle and Bergstrom 1992).
- Guidelines or procedures for performing benefits transfers is needed (Boyle and Bergstrom 1992; Smith 1992).
- Multi-site models need to be estimated for a resource or activity in order to better account for changes in characteristics and affects of substitutes across different sites (Desvousges, Naughton, and Parsons 1992).

- Compare multi-site models of the same structure from different regions, thus enabling the identification of important explanatory variables (Desvousges, Naughton, and Parsons 1992).
- Models need to incorporate policy relevant quality variables, increasing the applicability of original studies to future issues (Desvousges, Naughton, and Parsons 1992).
- Multi-site models need to experiment with readily accessible and regionally relevant explanatory variables (Desvousges, Naughton, and Parsons 1992).
- General population surveys are needed to investigate how people conceptualize the environmental resources involved in their consumption decisions (Smith 1992).
- Calibration functions need to be developed and transferred along with benefit estimates and or benefit functions to correct for benefits transfer bias (Feather and Hellerstein 1997).

This and the previous section identified several issues, problems, and recommendations for the future of benefits transfers. Due to the difficulties encountered in the application of benefits transfers empirically, these recommendations have been made with the hope of improving the process. All of this information will be important in accomplishing the current RPA benefits transfer project using meta-analytic techniques.

General findings are that point estimate transfers are riddled with bias. Benefit functions are preferred because they are more robust than point estimate transfers. Meta-analysis is a viable approach to investigate variability in benefit estimates due to methodological, site characteristics, and user population differences.

Other Recreation Meta-Analysis Studies

Two previous studies used meta-analytic techniques to assess recreation estimates from past research. The first was the Walsh, Johnson, and McKean (1988, 1989, 1992) work. They had 287 stated and revealed preference derived-benefit estimates from 120 studies. Their model included 21 explanatory variables across the three models developed – a pooled model, a stated preference model, and a revealed preference model. The variables used in defining their models primarily included methodological differences between the study estimates. Only two demographic variables were included, in which one of these was using Forest Service regions as a proxy for socioeconomic status. Two site characteristics proxy variables were also included. Other variables were for trend and recreation activity dummies.

Smith and Kaoru (1990) used meta-analytic techniques on 77 studies with approximately 400 benefit estimates in total. They estimated 8 different models with varying assumptions. Across the models, 21 explanatory variables were defined. These models also consisted primarily of methodological variables. No demographic variables were specified in the models. One site characteristic variable was included based on the type of the site. Other variables included a trend variable and recreation activity variables.

Preliminary Summary Statistics

The current database consists of approximately 700 benefit estimates from about 140 studies, with 64% using stated preference methods and 36% using revealed preference methods. This database includes the Walsh, Johnson, and McKean database and the additions made by Doug McNair, but does not include most fishing studies post-Walsh et al./McNair. The fishing data will be provided by another agency at a later date.

Over 100 variables were coded in the database, attempting to identify all potential sources of information on the individual studies including identifiers, methods used, assumptions made, site and user characteristics of the sample, and so on. A majority of the variables are for methodology used. These methods are quite apparent in the reporting of the study results. However, we hoped this project would substantially add to the state-of-the-art by more fully exploring population and site characteristics. Unfortunately, there are severe constraints in the database for this endeavor. For example, less than 3% of the 700+ observations reported average income of the sample used to estimate the reported value. Similar results were found for the reporting of education (<1%), age (3%), and gender (13%).

These preliminary results are in direct conflict with the third and fourth criteria identified for effective and efficient benefits transfer – reporting the relationships between population characteristics and site characteristics with benefit estimates. Recommendations 1 through 3 – better reporting, model specification, and reporting standards and values database – directly address solutions to these problems. While these constraints may not be insurmountable in themselves, when taken in combination with resource constraints, the more fully developed functions may not be feasible. In conclusion, not much has changed in the reporting of study results since the litany of papers were published in the 1992 *Water Resources Research*.

Bibliography

Atkinson, S. E., T. D. Crocker, and J. F. Shogren. 1992. "Bayesian exchangeability, benefit transfer, and research efficiency." *Water Resources Research* 28(3): 715-722.

Boyle, K. J., and J. C. Bergstrom. 1992. "Benefit transfer studies: Myths, pragmatism, and idealism." *Water Resources Research* 28(3): 657-663.

Brookshire, D. S., and H. R. Neill. 1992. "Benefit transfers: Conceptual and empirical issues." *Water Resources Research* 28(3): 651-655.

Desvousges, W. H., M. C. Naughton, and G. R. Parsons. 1992. "Benefit transfer: Conceptual problems in estimating water quality benefits using existing studies." *Water Resources Research* 28(3): 675-683.

Downing, M., and T. Ozuna, Jr. 1996. "Testing the reliability of the benefit function transfer approach." *Journal of Environmental Economics and Management* 30(3):316-322.

Eade, J. D. O., and D. Moran. 1996. "Spatial economic valuation: Benefits transfer using Geographical Information Systems." *Journal of Environmental Management* 48(2):97-110.

Feather, P., and D. Hellerstein. 1997. "Calibrating benefit function transfer to assess the Conservation Reserve Program." *American Journal of Agricultural Economics* 79(1):151-162.

Freeman, A. M., III. 1984. "On the tactics of benefit estimation under Executive Order 12291." In: V. K. Smith (ed.), <u>Environmental Policy Under Reagan's Executive Order: The Role of Benefit-Cost Analysis</u>, Chapel Hill, NC: The Univ. of North Carolina Press. pp. 167-186.

Kirchhoff, S., B. G. Colby, and J. T. LaFrance. 1997. "Evaluating the performance of benefit transfer: An empirical inquiry." *Journal of Environmental Economics and Management* 33(1):75-93.

Krupnick, A. J. 1993. "Benefit transfers and valuation of environmental improvements." *Resources* 110: 1-6.

Loomis, J. B. 1992. "The evolution of a more rigorous approach to benefit transfer: Benefit function transfer." *Water Resources Research* 28(3): 701-705.

Loomis, J., W. Provencher, and W. G. Brown. 1990. "Evaluating the transferability of regional recreation demand equations." In: R. L. Johnson and G. V. Johnson (eds.), Economic Valuation of Natural Resources: Issues, Theory, and Applications, Boulder, CO: Westview Press. pp. 205-217.

Luken, R. A., F. R. Johnson, and V. Kibler. 1992. "Benefits and costs of pulp and paper effluent controls under the Clean Water Act." *Water Resources Research* 28(3): 665-674.

McConnell, K. E. 1992. "Model building and judgment: Implications for benefit transfers with travel cost models." *Water Resources Research* 28(3): 695-700.

Parsons, G. R., and M. J. Kealy. 1994. "Benefits transfer in a random utility model of recreation." *Water Resources Research* 30(8):2477-2484.

Smith, V. K., and Y. Kaoru. 1990. "Signals or noise? Explaining the variation in recreation benefit estimates." *American Journal of Agricultural Economics* 72(2): 419-433.

Smith, V. Kerry. 1992. "On separating defensible benefit transfers from 'smoke and mirrors'." *Water Resources Research* 28(3): 685-694.

Sorg, C. F., and J. B. Loomis. 1984. "Empirical estimates of amenity forest values: A comparative review." *General Technical Report RM-107*, Rocky Mountain Forest and Range Experiment Station, USDA Forest Service.

Walsh, R. G., D. M. Johnson, and J. R. McKean. 1988. "Review of outdoor recreation demand studies with nonmarket benefit estimates, 1968-1988." *Technical Report No. 54*, Colorado Water Resources Research Institute, Colorado State University.

-----. 1989. "Issues in nonmarket valuation and policy application: A retrospective glance." *Western Journal of Agricultural Economics* 14(1): 178-188.

-----. 1992. "Benefit transfer of outdoor recreation demand studies, 1968-1988." *Water Resources Research* 28(3): 707-713.

ABSTRACT

WILLINGNESS TO PARTICIPATE IN ECOSYSTEM MANAGEMENT OF NIPF LANDS IN THE NORTHEAST

T.H. Stevens, R. Belkner, D. Kittredge, D. Dennis, C. Willis

Contingent valuation and conjoint analyses were used to examine forest landowner attitudes and preferences toward ecosystem management, the types of ecosystem management programs they would be willing to adopt and the likelihood of cooperation with neighbors to achieve ecosystem management. Results of two separate case studies suggest that the likelihood of undertaking ecosystem management was not reduced when collaborative management is required. Programs emphasizing wildlife habitat and protection of rare plant species were more likely to be accepted than programs focusing on timber harvests or recreation. A comparison of CV and conjoint results indicate that when the questions are the same in all respects, except for rating and pricing formats, median WTP estimates are different.

WILLINGNESS TO PARTICIPATE IN ECOSYSTEM MANAGEMENT OF NIPF LANDS IN THE NORTHEAST[a]

T. H. Stevens, R. Belkner, D. Kittredge, D. Dennis, C. Willis

Introduction

Ecosystem management (EM) is often defined as ecologically based, sustainable management that blends environmental and social values. Instead of focusing on commodity outputs, the ecosystem approach seeks to achieve desired future conditions, with outputs such as timber harvests, wildlife, and recreation opportunities occurring throughout the process (Stanley, 1995). Effective EM requires planning on broad spatial and temporal scales. And, particularly in the Northeastern United States, where the majority of forestland is held in relatively small parcels by nonindustrial private owners (NIPF), owner cooperation is essential in accomplishing EM objectives. As noted by Brunson, et al. (1996), "because most ecosystems do not conform with property lines, strategies that coordinate activities among multiple entities--including public agencies and private landowners--will be required in implementing ecosystem based management" (p.15).

Yet, very little is known about NIPF landowner attitudes and preferences toward EM, the types of EM programs they would be willing to adopt, or about the likelihood of cooperation with others to achieve ecosystem management. Moreover, debate continues about how this kind of information should be obtained. For example, contingent valuation is widely used for valuing environmental programs, but this approach is often viewed with skepticism. Alternatives to CV, such as conjoint analysis, CJ, have been explored, but few comparisons of CV and CJ analysis have been published (Boxall, et al, 1996).

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This paper presents two case studies that focus on these issues. The first study uses conjoint analysis to examine attitudes of randomly selected landowners about EM. The second study, which compares CV and conjoint techniques, focuses on landowners currently enrolled in a forest stewardship program.

Background

Although debate continues about the concept of ecosystem management, much of the literature suggests that compared to traditional approaches, ecosystem management emphasizes ecological principles while incorporating a wide range of societal values. Brunson, et al. (1996) argue that..."even if the concept as a whole is ultimately rejected, some aspects of it are likely to become part of the overall multiple-use tool kit" (p. 14). A potential problem, however, is that many NIPF owners may not be willing to participate in the types of collaborative efforts necessary for effective ecosystem management.

For example, Brunson, et al. (1996) surveyed NIPF owners about attitudes toward collaborative management in three regions: the Southeast, Midwest, and Interior West. When asked to respond to the statement that "public and private landowners should plan activities jointly because ecosystems and wildlife cross property boundaries", about 80% of all respondents rated this concept as "appropriate" or "highly appropriate" for public lands. But, only 23% of Utah and Southeast respondents and 14% of Midwest (Indiana) respondents said they would definitely be willing to plan activities jointly if their own land was involved. Most wanted to see a demonstration project before deciding whether to participate. Brunson, et al. (1996) conclude that....." many NIPF owners are deeply concerned about property rights, and this may make them less supportive of ecosystem management strategies that call for power-sharing among groups of landowners. However, wariness about property rights may be offset by NIPF owners deeply rooted beliefs about forest stewardship and the need to protect natural environments" (p. 20).

The following case studies focus on the attitudes of Massachusetts NIPF owner's about collaborative EM efforts and their WTP to undertake various ecosystem management activities.

Case Study 1

The first case study involved a 1995 survey of 1,250 randomly selected Franklin County, Massachusetts residents owning 10 acres or more of forestland. The survey instrument was designed and pretested using input from focus groups; a modified Dillman (1978) Total Design Method was used throughout.

Landowners were partitioned into two groups. Each received an identical questionnaire except that one group was asked about cooperative management options for a hypothetical set of adjacent privately owned parcels while the other was asked about the same options for a single equivalent parcel owned by the individual. Relevant portions of both surveys are presented in Appendix 1.

The conjoint method was used to elicit information about the probability that individuals would participate in EM programs. Respondents were asked to rate four alternative management scenarios, each of which consisted of a bundle of attributes or management activities, including cost, on a scale of 1 to 10, with 10 indicating scenarios, if any, the individual would definitely undertake, 1 representing scenarios the respondent would definitely not undertake, and if not sure, a rating of 2-9 indicated the likelihood that a scenario might be adopted (see Appendix 1).

Only those individuals who said they would definitely undertake each management scenario were counted as participating in EM. We assumed that each individual's decision to participate depended upon socio economic characteristics, such as age, education and income as well as five EM program attributes; extent of timber harvested, management cost, establishment of a recreational trail system, maintenance of apple trees and preservation of a rare species of ferns. Each attribute was assigned three different levels (e.g., harvest all, one-half or none of the timber; protect all, one-half or none of the ferns), and three management cost levels for each scenario were defined; $50, $250 , and $500 (see the Appendix), giving 243 possible combinations. An orthogonal array was then used to create the most succinct subset of all attribute combinations. The resulting 18 alternative management programs were then assigned to the 1,250 questionnaires in equal frequency.

The useable response rate to this survey was 61.3%. About 49% of respondents were 35 to 54 years old, and 74% had completed at least 1 year of college. The average respondent owned 66 acres of forestland, about 20% had filed a formal forest management plan, and 54% had owned their land less than 20 years.

In addition to the conjoint questions, all respondents were asked a series of questions about attitudes toward cooperating with neighbors on land management projects such as habitat improvement, timber harvesting, or recreational trail development. Only 18% agreed with the statement that "I would not consider cooperating with my neighbors on land management projects". Fifty-six percent said that they would agree to participate in and share the costs of occasional, specific management projects such as building walking trails or arranging for a timber sale with their neighbors. And, 28.6% said they would agree to "enter into a contractual agreement with neighboring landowners for a fixed period of time (e.g. 5-10 years) to hire one or several professionals to prepare and implement a management plan on all the properties collectively whereby all participating landowners would share equally in the benefits (e.g. timber sale revenues, recreational access, enhanced wildlife habitat) and the expenses (e.g. preparation of the plan, administration of timber sale, establishment of trails)."

The probability that landowners would undertake any of the EM programs presented in the survey was calculated using a binary logit model.

$$(1) \qquad E(Y) = \frac{1}{1+e^{-\alpha-\beta x}}$$

where Y equals 1 for programs that would definitely be undertaken by an individual (conjoint rating = 10) and Y equals 0 otherwise (conjoint rating = 1-9), x is a vector of the explanatory variables defined in Table 1 and α and β are estimated coefficients. Data from both survey types (individual management and cooperative management) were pooled and a dummy variable for survey type, T=1 if cooperative version, was included to test for the effect of cooperative management on the probability of program adoption.

Results of this analysis are reported in Table 2. The value of the coefficient for survey type, T, was negative but not statistically significant indicating that the probability of adopting cooperative alternatives was not different than for identical individual management alternatives, all else held constant. As expected, probability of program adoption increased with apple trees, ferns, trails, and harvests, but the effect of harvest was not statistically different from zero.[1] The probability of program acceptance declined with cost and preference for noncooperation (Q15A). However, the likelihood of participation was not statistically related to landowner education or age, parcel size, or management plan.

The probability that several different management programs would be undertaken by respondents is presented in Table 3. In interpreting these results, it is important to note that all variables not indicated in Table 3 were set at their mean value (see Table 1).

As expected, an increase in management cost reduces the likelihood that EM programs would be undertaken and if fewer positive attributes are involved (no maintenance of apple trees, for example) the probability that a program would definitely be undertaken is very small.

Referring to Tables 2 and 3, low cost EM programs which maintain apple trees and protect ferns had the highest likelihood of adoption (see program 4, Table 3).

To summarize, this case study shows that NIPF owners in Western Massachusetts are not reluctant to engage in collaborative management efforts and that ESM programs focusing on wildlife habitat and protection of rare species are more likely to be adopted than programs emphasizing timber harvests. However, as shown in Table 3, the probability of undertaking any of the ESM programs examined is relatively low.

Case Study 2

The second case study involved a mail survey of all 1116 Massachusetts landowners enrolled in the Forest Stewardship Program. This program is a voluntary, federally funded program that entitles participants to share the cost of improving their forest land with the federal government. Consequently, participants in this survey were generally very knowledgeable about forest land management activities.

Landowners were partitioned into four groups. One group received a dichotomous choice CV format containing cooperative EM management alternatives; the second group was given a CV format with private management alternatives; the third received a cooperative conjoint CJ question while the last group was given a private management CJ question format.

From the perspective of neoclassical economic theory, the CV and CJ formats should produce similar results, provided they are properly specified. Suppose that individual utility associated with EM can be expressed as a function of income, Y, and EM attributes such as water quality, wildlife habitat preserved, and cost. In dichotomous choice CV, individuals are asked to undertake an EM program that costs a predetermined amount, $N. The value of utility, observed by the researcher, when amount N is paid is:

(2) $$U_1 = U(D, Y\text{-}N) + e_1$$

where D is a vector of EM attributes and e is a random variable. The expected utility when $N is not paid is:

(3) $$U_0 = U(Y) + e_0$$

The individual is assumed to pay if, and only if:

(4) $$U_1 \geq U_0$$

The willingness to pay probability can then be written as:

(5) $$P_r = G(dV)$$

where G is the probability function for the random component of utility and dV is the expected utility difference:

(6) $$dV = U_1\text{-}U_0$$

If utility is assumed to be liner, additive, and separable with respect to income and EM attributes, dV is given by:

(7) $$dV = U(D) + U(\text{-}N) + e_1\text{-}e_0$$

Assuming a logit probability function for G in equation (5), the WTP probability is:

(8) $$P_r = (1+e^{-dV})^{-1}$$

Median WTP can then be estimated by calculating the value of N, N*, for which dV=0, i.e., at the point of indifference there is a 50 percent chance that the individual would pay amount N*.

Following Roe, et al. (1996) a CJ format which is conceptually consistent with the dichotomous choice CV format (eq. 8) can be derived by asking individuals to rate the current situation without an EM program as given by (3) and a set of alternative EM programs, (eq. 2). It is implicitly assumed that:

(9) $$R_1 = h(U_1), \text{ and } R_0 = h(U_0)$$

where R_1 and R_0 are individual ratings and h is a transformation function. Utility difference, dV, is then given by the ratings difference R_1- R_0:

(10) $$dV = R_1 - R_0 = U(D) + U(-N) + e_1 - e_0$$

where (10) is the same as (7).

If individuals are asked, for example, to rate programs, including the status-quo, on a scale of 1 to 10 with 10 indicating programs, if any, they would **definitely undertake**, a binary response model which is identical to the dichotomous choice CV model, (Eq. 8), is obtained.

It is important to note that the CJ model set forth in (9)- (10) differs from the traditional CJ format in that the dependent variable in (10) is the ratings difference from the status quo and independent variables are changes in program attributes from the status quo. Also, as shown by Roe, et al. (1996), this specification provides estimates of Hicksian surplus (also see McKenzie, 1990, 1993).[2]

Although very few CJ, CV empirical comparisons have been published, the evidence suggests substantial differences. One reason is that CV respondents are typically presented with far fewer substitutes as compared with CJ respondents and consequently CV results may be biased upward (Boxall, et al., 1996). On the other hand, CV may create incentives for respondents to not state their true value (Magat, et al., 1988). Moreover, Irwin, et al. (1993) argue that CV and CJ (or choice) results will generally differ because respondents usually do not have well defined monetary values for environmental commodities. For example... "when dollars are an available (recomprisable) attribute of an object, they carry more weight or influence in determining an evaluative response that is also in dollars (e.g., willingness to pay, cash equivalent, selling price) than they do in determining a response that is not in dollars (e.g., a rating of value or a choice)."

In contrast to most previous studies, the CV and CJ questions used in this study were virtually identical; any differences in WTP should therefore be due solely to the different response formats. All respondents were asked to value (rate) the status quo (do nothing) and three EM alternatives which involved setting aside a portion of their land to create a buffer zone that would provide a wildlife corridor connecting two larger wildlife habitats; a state forest and town conservation land. Respondents were told that the buffer zone would also improve water quality downstream to maintain a wood turtle population located on the town conservation land (see Appendix 2).

Each EM option consisted of three attributes; acreage set aside for the buffer zone, increase in wood turtle population, and annual improvement and maintenance costs associated with the buffer zone. There were three possible values for buffer zone acreage, 5, 10 or 20 acres, three levels of increase in the wood turtle population, 0%, 10%, 25%, and three annual cost levels, $50, $100, $200. All possible combinations of these were generated and three EM programs, plus the status quo, were randomly assigned to each respondent.

The CV response rate was 67% and 42% of the CV surveys returned were fully completed. The CJ response rate was 56%, but 67% were completed giving the CJ format a higher completed return rate.

Of particular importance is that tests for pooling showed that the cooperative and private management versions of this survey could be combined. Consequently, we focus on comparison of WTP estimates derived from the CV and CJ formats.[3]

Four different econometric models were estimated; a dichotomous choice CV logit model, two CJ logit models, and a 'ratings difference' CJ model. An approximation of utility difference, dV, was used in each specification:

(11) $dV = a + b(D) + c(N) + d(F) + e$

where D is a vector of EM attributes (acres, turtles), N is the predetermined program cost, F is a set of taste and preference variables which differ among individuals, and a, b, c, and d are estimated coefficients.

The dependent variable in the first CJ model, CJ1, equals 1 if the respondent would definitely undertake an EM program (rating equal to 10), and 0 otherwise. The dependent variable in the second CJ model, CJ2, equals 1 if the respondent rated an EM program greater than the status quo and 0 otherwise. The third CJ model, CJ3, is a more traditional specification wherein equation (11) is estimated using the Tobit procedure. As shown by Roe, et al. (1996), WTP is derived from CJ3 by increasing the value of N until the point of indifference is reached, ($dV=0$).

Independent variables are presented in Table 4. The first variable, acres, is the amount of land respondents were asked to set aside for the purpose of EM. We expect a negative relationship between acres and WTP, all else held constant. The second variable, WTINC, is the percentage increase in the wood turtle population as a result of EM. A positive relationship between WTINC and WTP is expected. The cost variable is the monetary commitment incurred by respondents undertaking EM programs. As noted above, three annual cost levels were used; $50, $100, and $200. These amounts were determined by analyzing cost information provided by the Massachusetts Forestry Stewardship Council. Clearly, an increase in cost should decrease WTP. Three variables, age, income and environment were used to represent socioeconomic characteristics of respondents. The environment variable is a binary variable which takes a value of 1 if a respondent agreed with the statement, '*the environment should be given priority even if it hurts the economy*', and 0 otherwise.

Estimated CV and CJ model coefficients are presented in Table 5. With the exception of acres, estimated CV coefficients were of the expected sign. However, only two variables, increase in wood turtle population and costs were statistically significant. The CJ model results were much more robust; all coefficients had the expected sign and relative to the CV model, more variables were statistically significant.

Estimated WTP was derived from the CV, CJ1, and CJ2 models for the 'average' EM program by using equation (8). The value of dV in (8) was obtained by multiplying the mean value of all independent variables, except cost, by the appropriate estimated coefficients (Table 5). Median WTP was then derived by calculating the cost that yields a 0.5 payment probability (see equation 8). Mean WTP values were calculated by integrating over the $0 to $200 cost range. The CJ3 WTP estimate was derived by finding the value for N which sets dV in equation (11) equal to zero.

Results of these calculations are presented in Table 6. The confidence intervals reported in Table 6 were estimated using a bootstrapping method wherein 300 random observations, selected with replacement, were used to generate 1000 WTP estimates for each of the models.

As shown in Table 6, the CV and CJ1 model median WTP point estimates are quite different and the confidence intervals barely overlap at the 95% level. It is important to emphasize that from the perspective of economic theory and econometric technique, these models are virtually identical. The only difference is that while CV respondents were asked if they would pay a predetermined amount, CJ1 respondents were asked to rate each EM option on a scale of 1 to 10 with 10 indicating that they would definitely undertake EM. Given these results, we believe that response foormat (e.g., CV or CJ) may be more important than previously thought.

The CJ2 and CJ3 model WTP estimates are only indirectly comparable with the CV result.

This is because it is assumed that CJ2 respondents would undertake all EM programs that were rated above the status quo, and the CJ3 specification assumes that rating difference is a cardinal measure of respondent preferences (Roe, et al., 1996).

As expected, the CJ2 median value point estimate is greater than the CV estimate and this difference is statistically different at the 95% level. In other words, the CJ2 median WTP estimate is biased upward because it is implicitly assumed that all respondents who 'might' pay will, in fact, do so. It is important to note that the CJ3 model results should also be interpreted as an upper bound because this specification does not distinguish between those who are or are not actually in the market for the commodity being valued.

Estimated probabilities that respondents would undertake several different types of EM programs are presented in Table 7. The CJ1 model estimates are much lower than either the CV or CJ2 estimates. Also the CJ1 probabilities presented in Table 7 are quite similar to those found in the first survey (see Table 3). However, it is important to emphasize that from a policy perspective, the estimated probabilities are extremely sensitive to model specification.

Conclusions

A majority of the forest landowners who were selected at random expressed interest in ecosystem management. However, the estimated probability that EM programs would actually be adopted by this group was relatively low. On the other hand, the likelihood of undertaking EM programs involving cooperative management was not statistically less than that associated with private management. And, a case study of landowners already enrolled in a forest management incentive program indicated that the probability of adopting EM programs was not reduced when collaborative management is required. This finding is important since landowner cooperation is essential for successful EM of NIPF lands in the Northeast. It is also important to note that EM programs emphasizing wildlife habitat and protection of rare plant species were more likely to be accepted than programs focusing on timber harvests and recreation (trail development).

Our comparison of CV and CJ techniques for estimating willingness to participate in EM suggests that when CV and CJ questions are the same, except for rating and pricing format, median WTP point estimates are different. Moreover, since most previous CJ studies have essentially counted 'maybe' responses as 'yes' responses, we believe that CJ WTP estimates have often been biased upwards.

Much more research comparing the CV and CJ techniques is needed. Although the CJ approach seems to offer several conceptual advantages relative to CV, CJ is very sensitive to model specification and results must therefore be carefully interpreted.

Table 1. Explanatory Variables Used in Conjoint Analysis: Survey 1

Variable	Definition	Mean	Standard Deviation
Age	Age of owner in years	54.2	13.0
Educ	Education level of owner (1-6) categories	3.93	1.53
Income	Household income level (1-6) categories	3.34	1.41
T	Survey version Dummy (1 if coop)	.50	.50
Apples	% of apple trees maintained (0, .5, 1)	.52	.40
Ferns	% of acres of rare ferns saved (0, .5, 1)	.50	.40
Trails	% of trail system improved (0, .5, 1)	.51	.41
Harvest	% of timberland harvested (0, .5, 1)	.47	.39
Cost	Net cost to landowner ($50, $250, $500)	263	184
Acres 1	Acres timberland owned	66	101
Plan	Dummy variable = 1 if management plan	.20	.40
Q15A	Scale variable for attitude about cooperation (1-5) with 5 indicating landowner is strongly opposed to cooperation	2.39	1.27

Table 2. Logit Conjoint Model Results: Survey 1

Variable	Estimated Coefficient	Chi-Square P value
Intercept	-3.61**	.0001
T	-.24	.19
Apples	1.36**	.0001
Ferns	1.11**	.0001
Trails	.49*	.06
Harvest	.17	.51
Cost	-.00107**	.04
Age	-.007	.37
Educ	-.046	.50
Income	.124*	.07
Acres 1	-.0009	.42
Plan	-.066	.79
Q15A	-.199**	.02

Chi-square = 85.71**

n = 1681

* significant at 10% level

**significant at 5% level

Table 3. Calculated Probability of Program Adoption: Survey 1

Program [a]	Probability of Adoption
1. Cost = $50, all apple trees maintained	.133
2. Average Cost ($263), all apple trees maintained	.109
3. Average Cost ($263), no apple trees maintained	.031
4. $50 cost, all apple trees maintained and all ferns saved, no timber harvest, 30 year old owner	.227

[a] All independent variables not indicated are set at their mean value (see table 1).

Table 4
Independent Variables: Survey 2

Variable	Expected Sign	Mean	Standard Deviation	Range Low	Range High
ACRES	-	11.83	6.24	5	20
WTINC (% increase)	+	11.85	10	0	25
COSTS ($)	-	$117.96	$61.64	$50	$200
INCOME (thousand $)	+	$57.24	$31.75	$7.50	$102.50
AGE	+/-	57	14	25	99
ENVIRONMENT	+	.1299	.34	0	1

Table 5
Estimated CV and CJ Models: Survey 2

	CV	CJ		
	CV DC Logit	CJ 1[a] Logit	CJ 2[b] Logit	CJ 3[c] Tobit
INTERCEPT	-0.7382 (0.1254)	-2.1796** (0.0001)	1.3038** (0.0209)	2.4867** (0.0437)
ACRES	0.0093 (0.5068) [1.009]	-0.0133 (0.3961) [.987]	-0.0385** (0.0122) [.962]	-0.0797** (0.0180)
WTINC	.073147** (0.0001) [1.08]	.060634** (0.0001} [1.06]	.042823** (0.0001) [1.04]	0.1133** (0.0001)
COSTS	-0.0044** (0.0036) [.996]	-0.0040** (0.0168) [.996]	-0.0061** (0.0001) [.994]	-0.0154** (0.001)
AGE	-0.0022 (0.7421) [.998]	0.0011 (0.8844) [1.001]	-0.0089 (0.2125) [.991]	0.0104 (0.5044)
INCOME	0.00421 (0.1384) [1.000]	0.00649* (0.0563) [1.316]	0.00706** (0.0301) [1.305]	0.01362* (0.0557)
ENVIRONMENT	0.1929 (0.4547) [1.213]	0.2745 (0.3300) [1.000]	0.2665 (0.3443) [1.000]	0.9791 (0.1076)
Observations	581	692	692	504

The values reported in the ( ) are χ^2p values.
* Significant at the .10 level
** Significant at the .05 level
The values reported in the [ ] represent exp (β_j), the odds ratio statistic.
[a] Dependent variable is 1 if individual would definitely undertake the program; 0 otherwise. (i.e., Rating=10)
[b] Dependent variable is 1 if individual rated program greater than the status-quo; 0 otherwise.
[c] Dependent variable is rating difference from status-quo.

Table 6
Estimated Willingness To Pay Values: Survey 2

	CV	CJ		
	CV DC Logit	CJ 1[a] Logit	CJ 2[b] Logit	CJ 3[c] Tobit
Estimated Median Values	$86	$-287	$211	$285
Estimated Mean Values[e]	$86	$31	$116	---

		CV	CJ		
		CV DC Logit	CJ 1[a] Logit	CJ 2[b] Logit	CJ 3[c] Tobit
Confidence[d] Intervals For Median Values	Low	$53	-585	216	295
	High	$111	$55	228	308

[a] Dependent variable is 1 if rating is 10; 0 otherwise.

[b] Dependent variable is 1 if program rated above status-quo; 0 otherwise.

[c] Dependent variable is ratings difference from status-quo.

[d] The low and high ranges are 95% confidence intervals which were constructed by bootstrapping from the original data set.

[e] Mean value calculated over $0 to $200 cost range.

Table 7. Calculated Probability of Program Adoption: Survey 2

Program[a]	Probability of Adoption		
	CV	CJ1	CJ2
$100 Cost	.48	.18	.66
$200 Cost	.38	.13	.52
$100 cost 25% increase in wood turtle population	.71	.32	.78
$200 cost; no increase in turtle population; 20 acre buffer zone	.22	.06	.32

[a] All independent variables not indicated are set at their mean value (see Table 4).

References

Boxall, P. W. Adamowicz, J. Swait, M. Williams and J. Larviere. 1996. "A Comparison of Stated Preference Methods for Environmental Valuation". Ecological Economics 18:243-253.

Brunson, M.W., D. Yarrow, S. Roberts, D. Guynn and M. Kuhns. 1996. "Nonindustrial Private Forest Owners and Ecosystem Management". Journal of Forestry. June: 14-21.

Dillman, D. 1978. Mail and Telephone Surveys: The Total Design Method. John Wiley and Sons, New York.

Irwin, J.R., P. Slovic, S. Licktenstein and G. McClelland. 1993. "Preference Reversals and the Measurement of Environmental Values". Journal of Risk and Uncertainty. 6:5-18.

Magat, W.A. W. K. Viscusi and J. Huber. 1988. "Paired Comparison and Contingent Valuation Approaches to Morbidity Risk Valuation". Journal of Environmental Economics and Management. 15:395-411.

McKenzie, J. 1993. "A Comparison of Contingent Preference Models". American Journal of Agricultural Economics. 75:593-603.

McKenzie, J. 1990. "Conjoint Analysis of Deer Hunting". Northeastern Journal of Agricultural and Resource Economics. 19(21):109-117.

Roe, K. Boyle and M. Teisl. 1996. "Using Conjoint Analysis To Derive Estimates of Compensating Variation". Journal of Environmental Economics and Management. 31:145-159.

Stanley, T.R. 1995. "Ecosystem Management and The Arrogance of Humanism". Conservation Biology. 9(2):255-262.

Endnotes

1. Sensitivity to model specification was investigated by repeating this analysis with the dependent variable equaling 1 for programs which received a conjoint rating greater than or equal to 7, and 0 otherwise. Signs and magnitudes of key coefficients, such as apples, ferns, cost, and harvests were very similar to those reported in Table 2.

2. The traditional conjoint model involves estimating the following relationship between ratings and program attributes:

$$U_i = R_i = V(Z^K) + P_z = b_0 P_z + b_1 Z_1^1 + \ldots b_n Z_n^1 + e_i$$

where U_i is individual i's utility for an attribute bundle; R_i is the individual's rating, $V(\cdot)$ is the non-stochastic component of the utility function, Z^K is a vector of attribute levels, P_z is the price for the attribute bundle Z, and b is the marginal utility or weight associated with each attribute.

Setting the total differential of (a) to the point of indifference and solving:

$$dU_i = b_0 dP_z + b_1 dZ_1^1 + \ldots = 0$$

yields marginal rates of substitution for the attributes Z_1^1. Since a price attribute, P_z, is included, the marginal utilities of all attributes can be rescaled into dollars, and willingness to pay for each attribute may be derived:

$$dP_z = -b_1 dZ_1^1 / b_0 \quad \text{or}$$

$$dP_z / dZ_1^1 = -b_1 / b_0$$

Another important aspect of the ratings difference CJ model specification is that in the traditional specification different respondents tend to center on different ranges of the ratings scale. Roe, et al., argue that this problem is avoided by using the status quo rating as a common anchoring point.

3. A 't' test was used to test for differences in socioeconomic characteristics of CV and CJ respondents. The null hypothesis that the two groups are the same was not rejected for age, but was rejected for income. The mean income of CJ respondents (\$59,844) was statistically different than the mean income of CV respondents (\$55,151).

APPENDIX 1

14. Please consider the hypothetical situation shown in Figure 1 in which your forest land is adjacent to two other parcels. Suppose you own property number 1 and that you are asked to consider cooperating with your neighbors for the purpose of managing your forest land as part of a larger unit. Four cooperative arrangements are presented on the next two pages, each of which is a set of activities that can be implemented on forest land in Franklin County, Massachusetts. Each arrangement has a net cost based on possible income from the sale of timber and expenses associated with other management activities. Please consider and compare the cooperative arrangements presented and indicate how you would rate each on a scale of 1 to 10. Please use 10 for arrangements, if any, that you would definitely undertake. Use 1 for arrangements, if any, that you would definitely not undertake. If you are not sure, use 2 through 9 to indicate how likely you would be to enter into each arrangement.

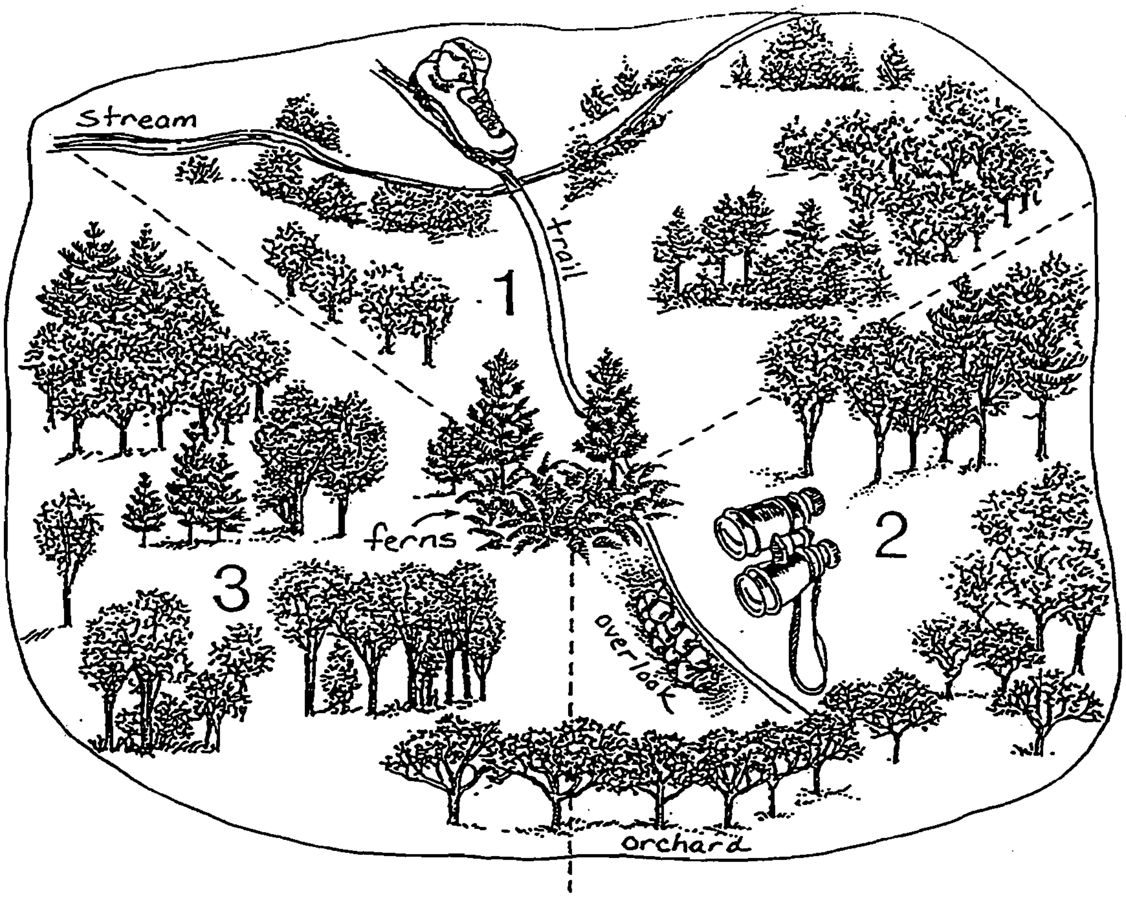


FIGURE 1

Arrangement A

- Agree to maintain of the apple trees shown on Figure 1 which benefit wildlife. Maintenance cost, if any would be shared equally with your neighbors.

- Agree to protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.

- Agree to improve of the trail network shown in Figure 1. The cost of improvements, if any, would be shared equally and would include the cost of building a footbridge over the stream and clearing scenic vistas.

- Agree to harvest timber from of the lands shown on Figure 1. Costs and revenues, if any, would be distributed to each landowner in proportion to a professional forester's estimate of value coming from each parcel. Any harvest would be selective, designed to remove poorly formed trees and leave some of high quality; 25-30% of all trees would be removed.

- This option would have a <u>net cost</u> to you of $. *RATING (1-10):* ________

Arrangement B

- Agree to maintain of the apple trees shown on Figure 1 which benefit wildlife. Maintenance cost, if any would be shared equally with your neighbors.

- Agree to protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.

- Agree to improve of the trail network shown in Figure 1. The cost of improvements, if any, would be shared equally and would include the cost of building a footbridge over the stream and clearing scenic vistas.

- Agree to harvest timber from of the lands shown on Figure 1. Costs and revenues, if any, would be distributed to each landowner in proportion to a professional forester's estimate of value coming from each parcel. Any harvest would be selective, designed to remove poorly formed trees and leave some of high quality; 25-30% of all trees would be removed.

- This option would have a <u>net cost</u> to you of $ *RATING (1-10):* ________

PLEASE CONTINUE TO THE TOP OF THE NEXT PAGE.

Arrangement C

- Agree to maintain of the apple trees shown on Figure 1 which benefit wildlife. Maintenance cost, if any would be shared equally with your neighbors.
- Agree to protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Agree to improve of the trail network shown in Figure 1. The cost of improvements, if any, would be shared equally and would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Agree to harvest timber from of the lands shown on Figure 1. Costs and revenues, if any, would be distributed to each landowner in proportion to a professional forester's estimate of value coming from each parcel. Any harvest would be selective, designed to remove poorly formed trees and leave some of high quality; 25-30% of all trees would be removed.
- This option would have a net cost to you of $. *RATING (1-10):* ________

Arrangement D

- Agree to maintain of the apple trees shown on Figure 1 which benefit wildlife. Maintenance cost, if any would be shared equally with your neighbors.
- Agree to protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Agree to improve of the trail network shown in Figure 1. The cost of improvements, if any, would be shared equally and would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Agree to harvest timber from of the lands shown on Figure 1. Costs and revenues, if any, would be distributed to each landowner in proportion to a professional forester's estimate of value coming from each parcel. Any harvest would be selective, designed to remove poorly formed trees and leave some of high quality; 25-30% of all trees would be removed.
- This option would have a net cost to you of $. *RATING (1-10):* ________

Individual Management Version

14. Please consider the hypothetical situation in which you own the forestland shown in Figure 1. Four management options are presented on the next page, each of which is a set of activities that can be implemented on your land in Franklin County, Massachusetts. Each arrangement has a net cost based on possible income from the sale of timber and expenses associated with other management activities. Please consider and compare the arrangements presented and indicate how you would rate each on a scale of 1 to 10. Please use 10 for arrangements, if any, that you would definitely undertake. Use 1 for arrangements, if any, that you would definitely not undertake. If you are not sure, use 2 through 9 to indicate how likely you would be to undertake each option.

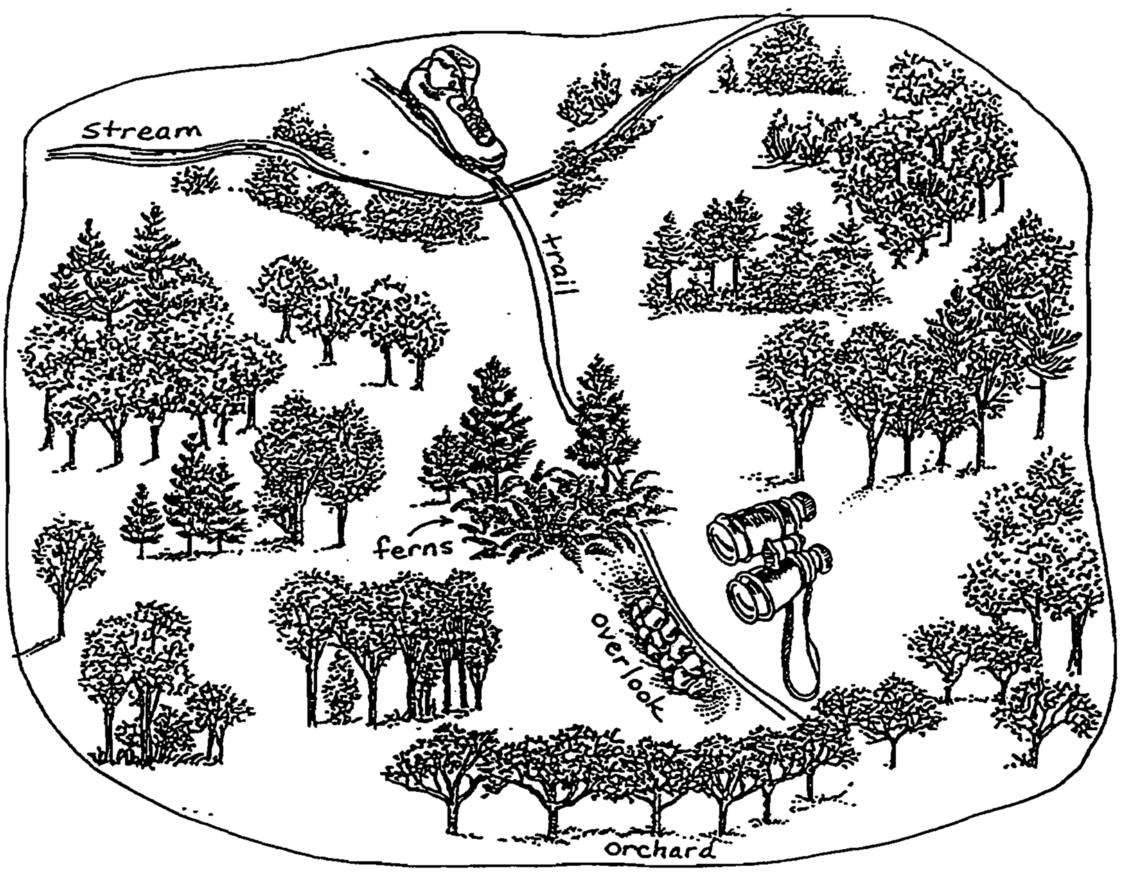


FIGURE 1

Option A

- Maintain of the apple trees shown on Figure 1 which benefit wildlife.
- Protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Improve of the trail network shown in Figure 1. Improvements, if any, would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Harvest timber from of the lands shown on Figure 1. Any harvest would be selective, designed to remove poorly formed and leave some high quality trees; 25-30% of all trees would be removed.
- This option would have a <u>net cost</u> to you of $

RATING (1-10): _______

Option B

- Maintain of the apple trees shown on Figure 1 which benefit wildlife.
- Protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Improve of the trail network shown in Figure 1. Improvements, if any, would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Harvest timber from of the lands shown on Figure 1. Any harvest would be selective, designed to remove poorly formed and leave some high quality trees; 25-30% of all trees would be removed.
- This option would have a <u>net cost</u> to you of $

RATING (1-10): _______

PLEASE CONTINUE TO THE TOP OF THE NEXT PAGE.

Option C

- Maintain of the apple trees shown on Figure 1 which benefit wildlife.
- Protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Improve of the trail network shown in Figure 1. Improvements, if any, would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Harvest timber from of the lands shown on Figure 1. Any harvest would be selective, designed to remove poorly formed and leave some high quality trees; 25-30% of all trees would be removed.
- This option would have a <u>net cost</u> to you of $.

RATING (1-10): ________

Option D

- Maintain of the apple trees shown on Figure 1 which benefit wildlife.
- Protect of the acres containing a rare species of fern shown in Figure 1 by not harvesting timber in this area or otherwise disturbing the ferns.
- Improve of the trail network shown in Figure 1. Improvements, if any, would include the cost of building a footbridge over the stream and clearing scenic vistas.
- Harvest timber from of the lands shown on Figure 1. Any harvest would be selective, designed to remove poorly formed and leave some high quality trees; 25-30% of all trees would be removed.
- This option would have a <u>net cost</u> to you of $.

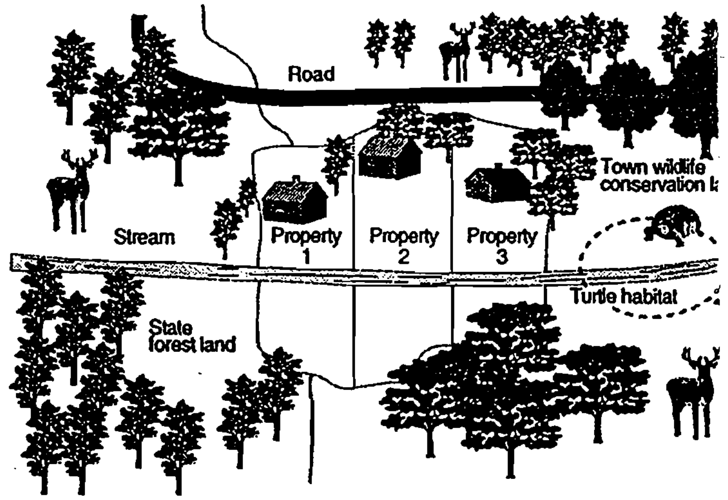
RATING (1-10): ________

APPENDIX 2

a. Conjoint Cooperative

b. CV Cooperative

Please consider the situation shown below. Suppose that you own and reside on pr erty number 2 which is adjacent to two other privately owned forested parcels. Each foreste parcel contains about 50 acres. 200 acres of state forest land is adjacent to one end of the forested parcels. Adjacent to the opposite end of the forest parcels is 600 acres of town wildlife conservation land. A stream runs through all five parcels of land. All land next to this stream is forested but is not suitable for any other land use, such as housing development. Wood turtles exist downstream on the town wildlife conservation land. <u>It is important to view the five separate parcels as one regional ecosystem, where the environmental functions of each parcel are interconnected. That is, land management decisions on one parcel of land impact envir mental functions on the surrounding parcels</u>.



Suppose that you are asked to cooperate with your neighbors for the purpose of m aging your land as part of a larger unit. Specifically, you are asked to agree to set aside, impro and maintain a buffer zone on each side of the stream. Improvements include planting of shru along the stream bank to reduce damage from runoff and sediment to downstream areas. <u>Fi decisions about improvements and the cost of all improvements will be shared equally with y neighbors</u>. This buffer zone creates a natural wildlife corridor; it connects the two larger parc of wildlife habitat, the state forest and town wildlife conservation land. The buffer zone a improves water quality downstream which is important for maintaining the wood turtle popu tion located on the town wildlife conservation land.

Please consider the following alternatives, each of which consists of several attribute Please consider and compare <u>all</u> the alternatives presented and <u>then</u> indicate how y would rate each on a scale of 1 to 10. Use 10 for alternatives, if any, that you WOUL DEFINITELY undertake. Use 1 for alternatives, if any, that you WOULD DEFINITELY N undertake. If you are not sure use 2 through 9 to indicate how likely you would be to ent into each alternative arrangement.

A13A1

You agree to set aside 10 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 0 % a13a2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 50 per year. This cost will also be incurred by your neighbors.

a13a3 **Alternative A**

Your rating for alternative A is ________ (scale 1 to 10)

a13a c5

a13b1

You agree to set aside 20 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 0 % a13b2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 200 per year. This cost will also be incurred by your neighbors

a13b3 **Alternative B**

Your rating for alternative B is ________ (scale 1 to 10)

a13bCJ

Do nothing

No buffer zone

No increase in wood turtle population

No additional improvement or maintenace costs

Alternative C

Your rating for alternative C is ________ (scale 1 to 10)

a13cCJ

a13d1

You agree to set aside 5 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 10 % a13d2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 200 per year. This cost will also be incurred by your neighbors

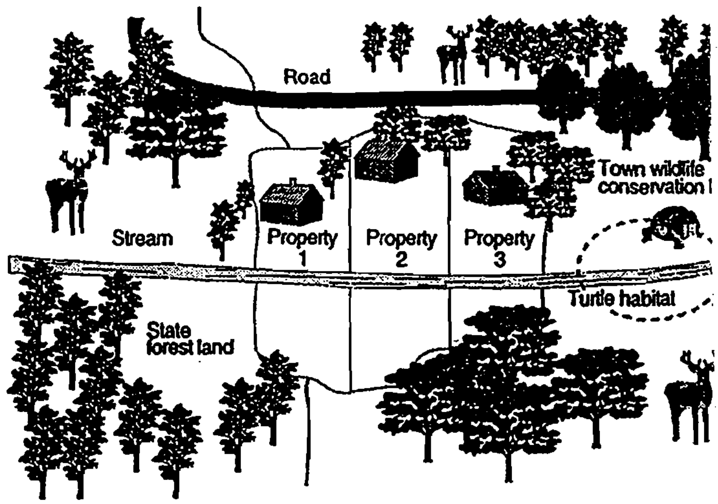
a13d3 **Alternative D**

Your rating for alternative D is ________ (scale 1 to 10)

a13dCJ

Please refold and complete section 3 on back

Please consider the situation shown below. Suppose that you own and reside c property number 2 which is adjacent to two other privately owned forested parcels. Ea forested parcel contains about 50 acres. 200 acres of state forest land is adjacent to one end of the forested parcels. Adjacent to the opposite end of the forest parcels is 600 acres of town wildlife conservation land. A stream runs through all five parcels of land. All land next to this stream is forested but is not suitable for any other land use, such as housing development. Wood turtles exist downstream on the town wildlife conservation land. It is important to view the five separate parcels as one regional ecosystem, where the environmental functions of each parcel are interconnected. That is, land management decisions on one parcel of l impact environmental functions on the surrounding parcels.



Suppose that you are asked to cooperate with your neighbors for the purpos managing your land as part of a larger unit. Specifically, you are asked to agree to aside, improve, and maintain a buffer zone on each side of the stream. Improvem include planting of shrubs along the stream bank to reduce damage from runoff and s ment to downstream areas. Final decisions about improvements and the cost c improvements will be shared equally with your neighbors. This buffer zone creates a ural wildlife corridor; it connects the two larger parcels of wildlife habitat, the state forest town wildlife conservation land. The buffer zone also improves water quality downstr which is important for maintaining the wood turtle population located on the town wi conservation land.

Please consider and compare all the alternatives presented and then indicate which al natives, if any, you would definitely undertake.

a13a1

You agree to set aside __5__ acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by __25__% a13a2

Your share of improvement and maintenence costs associated with the buffer zone will be $__56__ per year. This cost will also be incurred by your neighbors.

a13a3 **Alternative A**

Would you definitely undertake alternative A

___Yes

___No a13a

a13b1

You agree to set aside 10 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 25 % a13b2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 50 per year. This cost will also be incurred by your neighbors

a13b3

Alternative B

Would you definitely undertake alternative B

___Yes

___No a13bcv

Do nothing

No buffer zone
No increase in wood turtle population
No additional improvement or maintenace costs

Alternative C

Would you definitely undertake alternative C

___Yes

___No a13ccv

a13d1

You agree to set aside 5 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 10 % a13d2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 100 per year. This cost will also be incurred by your neighbors

a13d3

Alternative D

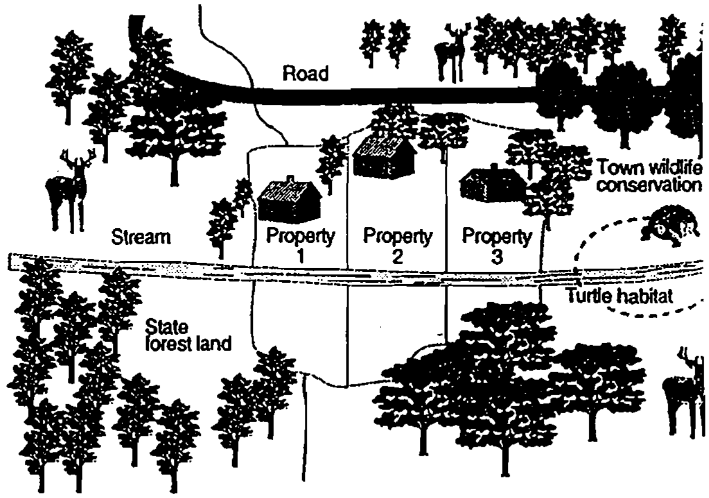
Would you definitely undertake alternative D

___Yes

___No a13dcv

Please refold and complete section 3 on back

Please consider the situation shown below. Suppose that you own and reside on p erty number 2 which is adjacent to two other privately owned forested parcels. Each forest parcel contains about 50 acres. 200 acres of state forest land is adjacent to one end of the forested parcels. Adjacent to the opposite end of the forest parcels is 600 acres of town wildlife conservation land. A stream runs through all five parcels of land. All land next to this stream is forested but is not suitable for any other land use, such as housing development. Wood turtles exist downstream on the town wildlife conservation land. <u>It is important to view the five separate parcels as one regional ecosystem, where the environmental functions of each parce are interconnected. That is, land management decisions on one parcel of land impact envi mental functions on the surrounding parcels</u>.



Suppose that you are asked to cooperate with your neighbors for the purpose of m aging your land as part of a larger unit. Specifically, you are asked to agree to set aside, impr and maintain a buffer zone on each side of the stream. Improvements include planting of shr along the stream bank to reduce damage from runoff and sediment to downstream areas. <u>E decisions about improvements and the cost of all improvements will be shared equally with y neighbors</u>. This buffer zone creates a natural wildlife corridor; it connects the two larger par of wildlife habitat, the state forest and town wildlife conservation land. The buffer zone a improves water quality downstream which is important for maintaining the wood turtle pop tion located on the town wildlife conservation land.

Please consider the following alternatives, each of which consists of several attribut Please consider and compare <u>all</u> the alternatives presented and <u>then</u> indicate how y would rate each on a scale of 1 to 10. Use 10 for alternatives, if any, that you WOU DEFINITELY undertake. Use 1 for alternatives, if any, that you WOULD DEFINITELY N undertake. If you are not sure use 2 through 9 to indicate how likely you would be to er into each alternative arrangement.

A13a1

You agree to set aside 10 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 0 % a13a2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 50 per year. This cost will also be incurred by your neighbors

a13a3 **Alternative A**

Your rating for alternative A is ________ (scale 1 to 10)

a13a5

a13b1

You agree to set aside 20 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 0 % a13b2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 200 per year. This cost will also be incurred by your neighbors

a13b3 **Alternative B**

Your rating for alternative B is ________ (scale 1 to 10)

a13bCJ

Do nothing

No buffer zone
No increase in wood turtle population
No additional improvement or maintenace costs

Alternative C

Your rating for alternative C is ________ (scale 1 to 10)

a13cCJ

a13d1

You agree to set aside 5 acres of your property for the buffer zone. Limited timber management, including harvest, could continue in this zone, but please assume that this land is not suitable for other uses, such as housing development. Population of the wood turtle will increase by 10 % a13d2

Your share of improvement and maintenence costs associated with the buffer zone will be $ 200 per year. This cost will also be incurred by your neighbors

a13d3 **Alternative D**

Your rating for alternative D is ________ (scale 1 to 10)

a13dCJ

Please refold and complete section 3 on back

DRAFT

Identifying Symbolic Effects in Contingent Choice Surveys: A Case Study of the Peconic Estuary System

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DRAFT

Extracting Symbolic Values from Contingent Choice Surveys: A Case Study of the Peconic Estuary System

Abstract

Techniques for correcting biases in CV surveys generally focus on *ex ante* methods for identifying bias during the survey development process, and adapting the survey in an attempt to circumvent the difficulties. Although this is undoubtedly the most effective means of correcting survey biases, it is generally not possible to anticipate and correct every potential source of bias before the fact. We may also be able to correct biases that are identified after the fact, during the process of data analysis. This paper discusses an attempt to identify and correct for symbolic effects in a survey undertaken to identify public values and priorities for important natural resources of the Peconic Estuary, under the National Estuary Program. The paper discusses the development of the survey, and presents a comparison of different estimation models, including two models that were developed to identify and isolate symbolic components of stated willingness to pay. Correcting for this symbolic component of willingness to pay results in a decline in resource values by 26% to 46%, depending upon the amenity and the model used.

I. Introduction

Contingent valuation and its variants are increasingly used as policy tools to measure values of natural amenities (see, for example, Carson). While the method has shown considerable potential (Carson, Flores and Meade, 1997) and has been viewed as a useful tool for measuring values for guiding public policy (e.g., Arrow et al, 1993; Ohio versus US Department of the Interior, 1989; U.S. Water Resources Council, 1983), various biases have also been identified (Arrow et al, 1993; Mitchell and Carson, 1989; Hausman, 1993). In order to minimize these biases, recommended procedures for CV surveys include a rigorous development process with focus groups and cognitive interviews. (e.g., Mitchell and Carson, 1989; Desvouges and Smith, 1984) Yet, measuring values of complex environmental amenities is an inherently difficult task, and the potential for bias remains in spite of the most rigorous development process. Various tests for identifying bias *ex post* in survey data have been proposed, include scope tests (e.g., Desvouges et al., 1993) and additivity tests (Diamond et al, 1993).

This paper attempts to go a step further by isolating and extracting one source of symbolic effects *ex post*, and calculating dollar values that are net this source of effect. We identify symbolic values by inferring "... when respondents react to an amenity's general symbolic meaning instead of to the *specific levels of provision described*." (Mitchell and Carson, 1989, P. 250, emphasis added). We use two models to isolate this component of stated willingness to pay, and calculate resource values net of the symbolic component. The paper is organized as follows. The Section II provides some information on the study, including background information on the area and the survey development and implementation processes. Section

III describes the models and the results. Section IV contains the summary and conclusions of the study.

II. Study Background

This paper describes a survey undertaken as part of a comprehensive economic value assessment undertaken for the Peconic Estuary Program (PEP), one of 28 estuary studies programs funded to date under the National Estuary Program. The Peconic Estuary system (see Figure 1), located at the eastern end of Long Island, New York, comprises over 100 bays, harbors, embayments, and tributaries. The Estuary's watershed drains a land area of around 110,000 acres (Suffolk County Department of Health Services 1994). The area surrounding the Estuary is a major recreation area for residents of New York City and western Long Island, and has a large seasonal population. Its economy includes a significant resource-based component, with important economic sectors associated with agriculture, fishing, wine making, tourism and recreation, and services related to second homes (East End Economic and Environmental Task Force 1994; Economic Analysis, Inc. 1995).

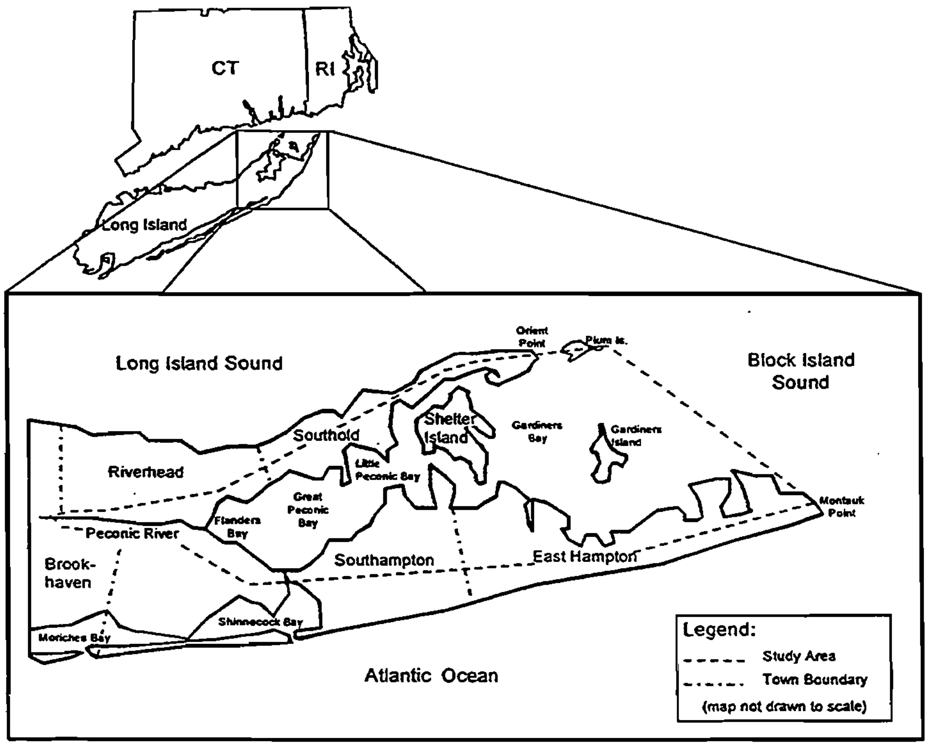
The East End's natural beauty and rural character, combined with its proximity to a large population center and the growing trend of telecommuting, have led to strong development pressures. Although environmental quality of the Estuary has historically been good to excellent, rapid development and associated environmental problems threaten the Estuary's waters and other amenities such as wetlands, farmland and other open space; commercial and recreational fish stocks; aquifers; and fish and wildlife habitat (Suffolk County Department of Health Services 1992).

This study is one part of a multidisciplinary process leading to the creation of a Management Plan for the Estuary. Natural science and other technical studies are simultaneously being carried out to determine the causes of, and potential solutions to Estuary's important environmental problems. Information on public preferences and economic values from this study will be combined with other components of the economic assessment, and with the science and technical information, in order to prioritize proposed actions for the final Management Plan.

Survey Development and Implementation

The primary goal of the survey was to elicit public priorities and values for natural resources of the Peconic Estuary that might be affected by preservation and restoration actions. The first stage of survey development consisted of meetings with the PEP Management Committee, the PEP Citizens' Advisory Committee, and knowledgeable representatives of stakeholder groups. This was followed by informal interviews of the public, focus groups, a brief preliminary survey, and pretesting of successive draft survey instruments. In addition, ten in-depth personal interviews were conducted in order to learn more about people's environmental attitudes and motivations for supporting different types of actions. Thus, the

Figure 1 - Map of Peconic Estuary System



final survey instrument was based on the needs of the Management Committee, combined with input from representatives of stakeholder groups and the general public.

Because specific actions and their results were not defined at the time the survey was developed and implemented, the survey was designed to elicit public preferences for generic natural resources that were most important to the public, and would most likely be affected by restoration and preservation programs. Therefore, the survey identifies public support for many different combinations of improvements in resources that might be achieved by management actions. This allows for assessment of programs, either by ranking or valuation of benefits, that affect any combination of the natural resources evaluated in the survey. Of course, a full assessment of management actions would require an analysis of the associated costs, including possible non-monetary costs of an action. For example, a program that limits fertilizer applications on lawns or that places limits on recreational boating might have non-monetary impacts that would need to be considered separately.

The contingent choice format, rather than contingent valuation, was selected for several reasons. Based on discussions following survey pretests in focus groups, the contingent choice framework seemed to be the most effective means of eliciting the required information. In the final survey, respondents were asked to select from a set of three hypothetical choices: no new action or one of two enhancement/protection programs. Each option was described by different levels of resulting natural resources, and the annual cost to each household, with the levels depicted in words as well as graphically. In early focus groups, only the two programs were presented, but focus groups participants indicated that a "no new action" option should be added, for two reasons. First, it allowed people to express a preference for no action if they did not support either program. Second, it provided a baseline from which to judge the benefits of each of the programs.

Based on concerns expressed by participants in focus groups and natural resources identified as important by the Technical Advisory Committee, five natural resources were selected to be included in the survey: farmland, undeveloped land, wetlands, safe shellfishing areas, and eelgrass. Based on the results of focus groups and past research (Mazzotta and Opaluch 1995), each comparison was designed to include only three attributes: two of the five natural resources included in the survey and the cost of the hypothetical resource protection program. This simplifies the choice, so that it is more likely that choices will be based on considering and balancing all of the attributes rather than using a simplified decision rule.

The objective of the survey was to elicit relative values for improvements in natural resources above a baseline level. The baseline was defined as the level that would exist in the year 2020 if no new action is taken to preserve or enhance the resource, and was determined based on historical declines and the judgment of experts.

In the survey, respondents were presented with background information, which described the level of each resource in 1981, the current level, and projected levels in 2020 if no new actions are taken. Respondents were told that "trends indicate approximate conditions in 2020," in order to make it clear that these are not scientifically-based projections, but are, in some cases, merely extrapolations of past trends. These levels are shown in Table 1 and the survey background information is shown in Figure 2. Figure 3 presents the instructions and an example contingent choice question.

In the contingent choice questions, each resource was included at three different levels: the projected level for 2020 (the "no new action," or baseline scenario), and two levels associated with hypothetical programs that would preserve or enhance natural resources. In order to make the hypothetical context clear, survey respondents were told "The following programs are hypothetical. We are trying to learn which resources are most important to you and how much you would pay to protect them." The levels for each resource and for cost are shown in Table 2.

A total of sixty different comparisons were created, with five in each of twelve different survey booklets. The combinations of attributes and levels were selected using a method based on Addelman's fractional factorial design, which produces orthogonal arrays of attribute main effects (Addelman 1962a, 1962b; Addelman and Kempthorne, 1961). The Addleman approach was modified to exclude alternatives where one program clearly dominated the other, and a small number of high-cost, high resource protection scenarios were added.

In addition to the contingent choice questions, the survey also asked respondents about recreational uses of the area's waters; level of support for specific resource protection actions; concern and knowledge about Brown Tide; and demographic information. The survey was completed by 968 respondents in August 1995 in a variety of public places throughout the area surrounding the Peconic Estuary.

Of the 968 people who completed the survey, 897 (92.7%) answered at least one of the five contingent choice questions in each survey booklet; and of the 4,840 total possible choices, 4,307 (89%) were answered. Older respondents, and those with lower education and income levels, were slightly less likely to answer all of the choice questions. Of those who answered at least one choice question, 91.6 percent chose either Program A or Program B, rather than "No New Action," for more than half of their answers, and 79 percent chose an action for all of their answers. Only 27 people (3.0%) chose "No New Action" for all of the choice questions answered.

These results demonstrate the strong environmental concern expressed by survey respondents, but also may suggest symbolic bias to the extent that respondents chose to "take action to protect the environment," without considering the specific levels of resource protection. In our case symbolic effects associated with "taking action" will show up as a common factor to the two action alternatives that is independent of the associated levels of resource protection. Thus, our survey data should allow us to identify the extent to which choices are based on the level of resource protection provided by the programs, versus the symbolic desire to "take action", independent of the specific level of provision of the resource amenities. We interpret the latter as symbolic bias, consistent with the definition of symbolic bias in Mitchell and Carson (1989), quoted above. Below we discuss some methods that we employed to identify and control for this possible bias.

Table 1- Past, Present and Projected Natural Resource Levels

Natural Resource	1981 Level	1995 Level	Projected 2020 Level
Farmland[a]	13,500 acres	12,000 acres	9,000 acres
Undeveloped Land[a]	74,000 acres	66,000 acres	50,000 acres
Wetlands[b]	18,000 acres	16,000 acres	12,000 acres
Safe Shellfishing Areas[b]	28,000 acres	26,000 acres	25,000 acres
Eelgrass[b]	10,000 acres	9,000 acres	8,000 acres

a - Calculated based on Long Island Regional Planning Board 1981, and Suffolk County Department of Health Services 1992.

b - Calculated based on Suffolk County Department of Health Services 1992 and information provided by NY State DEC.

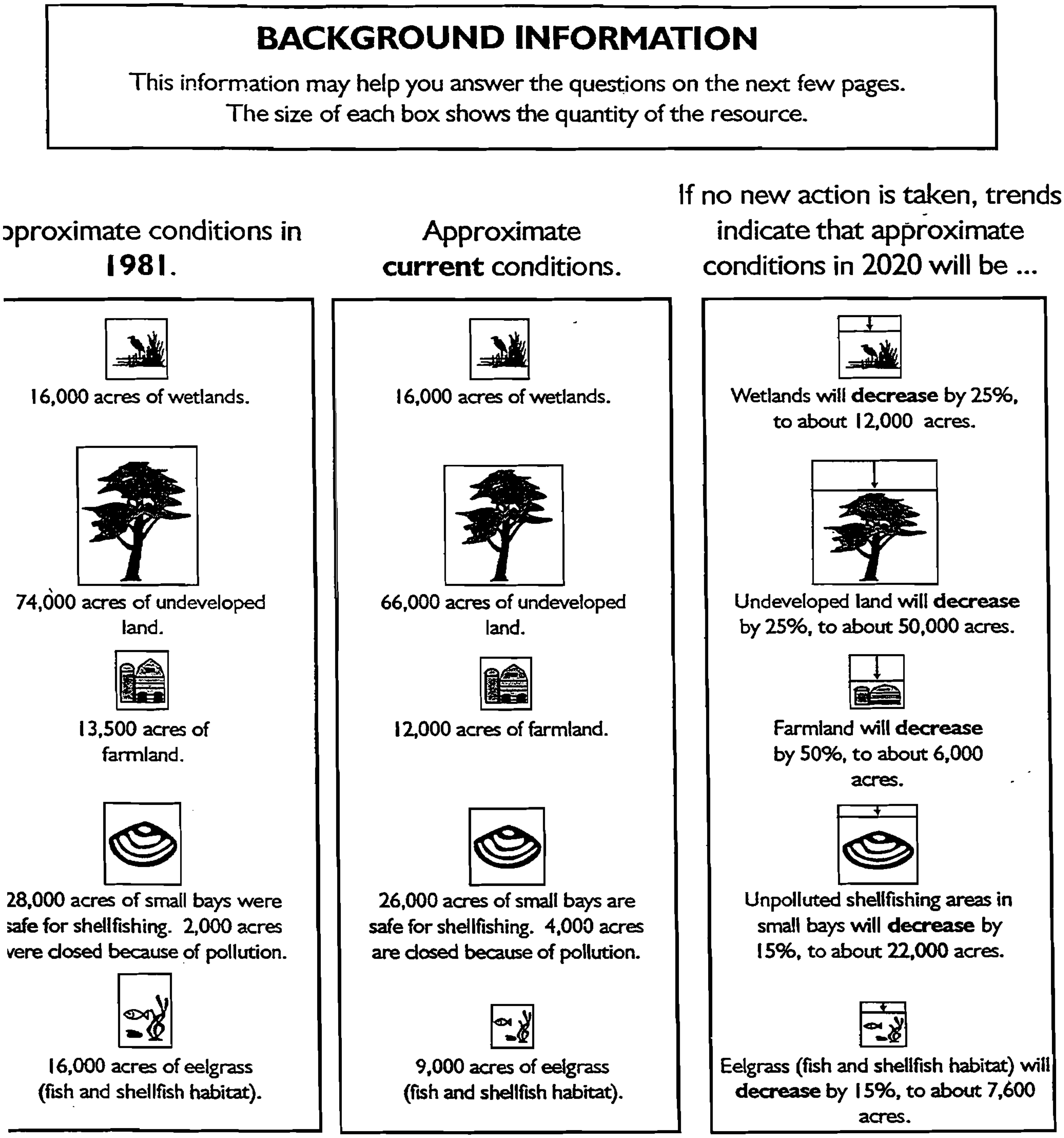


Figure 2. Survey Background Information

If you had to choose one of the 3 options below, which would you choose?

Circle Program A, Program B, or No New Action below.
(Do not compare these to programs on any other page.)

Projected Results for 2020:

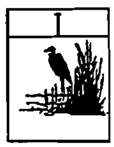
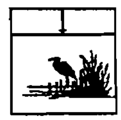
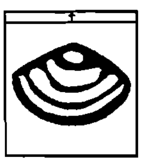
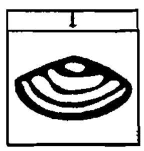
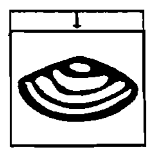
Program A	Program B	No New Action
		
Current wetlands will be preserved at16,000 acres.	Wetlands will **increase** by 15%, to about 18,000 acres.	Wetlands will **decrease** by 25%, to about 12,000 acres.
		
Unpolluted shellfishing areas in small bays will **increase** by 10%, to about 29,000 acres.	Unpolluted shellfishing areas in small bays will **decrease** by 15%, to about 22,000 acres.	Unpolluted shellfishing areas in small bays will **decrease** by 15%, to about 22,000 acres.
$300 per year	$100 per year	$0 per year
Cost to each East End household: $300 per year.	Cost to each East End household: $100 per year.	No new cost to each East End household.

Figure 3. Sample Survey Question

Table 2 - Levels of Natural Resources and Cost in the Survey

	Baseline Level in 2020 with No Action (% Change)	Level with Moderate Preservation or Restoration (% Change)	Level with High Preservation or Restoration (% Change)
Farmland	9,000 acres (-25%)	10,000 acres (-15%)	12,000 acres (current)
Undeveloped Land	50,000 acres (-25%)	59,000 acres (-10%)	63,000 acres (-5%)
Wetlands	12,000 acres (-25%)	16,000 acres (current)	17,500 acres (+10%)
Safe Shellfishing Areas	25,000 acres (-5%)	26,000 acres (current)	29,000 acres (+10%)
Eelgrass	8,000 acres (-10%)	9,000 acres (current)	11,000 acres (+25%)
Cost Levels	$0 $50	$100 $200	$300 $500

III. Model and Results

Based on the random utility model, relative values and priorities for the natural resources were estimated using the standard conditional logit method (Greene 1993; Maddala 1983), where

$$P_{ij} = \frac{\exp(\beta' z_j + \alpha' z_j w_i)}{\sum_k \exp(\beta' z_k + \alpha' z_k w_i)}. \qquad (1)$$

P_{ij} is the probability that individual i will select option j; z_j is a vector of attributes of the choice (e.g., the levels of natural resources and the cost), which may also vary across individuals; w_i is a vector of characteristics of the individual; and β and α are vectors of parameters of the model, estimated using maximum likelihood techniques.

Box-Cox tests were carried out to select an appropriate functional form and, based on these tests, the linear form was chosen for all subsequent estimation. The first model includes only the attributes of the choices—the levels of natural resources and cost. The results for this model are reported in Table 3. The model results indicate that the order of priorities for protection or enhancement of resources is as follows: farmland, eelgrass, wetlands, shellfish, and undeveloped land.

The results of the conditional logit model were compared to two alternative specifications designed to isolate symbolic values, as defined above. The first is the conditional logit model with two alternative-specific constants, one for the choice of an action versus no action, and one for Program B. Thus, the coefficient on the first constant term reflects factors other than the levels of attributes that affect the choice of an action versus the choice of no action. For example, respondents may be expressing a symbolic willingness to pay to "take action to protect the environment", as opposed to revealing incremental values for the "... specific levels of provision described" (Mitchell and Carson, 1989, P. 250). Here we are trying to measure specific amenity values, so we wish to exclude the component associated with the desire to "take action" *per se*. The coefficient on the second constant term reflects any difference in preference for Program B versus Program A that is unrelated to the levels of attributes of A and B, such as an order effect.

The second alternative specification is the nested logit model, which allows for correlations between the error terms of the two action alternatives. Such correlation implies that individuals view the two "action" alternatives as different from "no new action" in ways other than can be explained by the specific level of provision of resource amenities, which is represented by the explanatory variables. Again, this correlation may indicate a symbolic effect associated with "taking action to protect the environment", in addition to a quantitative assessment based on the specific level of resource protection provided.

Table 3 - Conditional Logit Base Model Results (N=4307)

Variable	Coefficient	Std. Error	T-Stat.
Farmland	.0005113	.000028	17.94
Undeveloped Land	.0001066	.000006	18.06
Wetlands	.0003361	.000017	20.04
Shellfish Areas	.0002334	.000019	12.38
Eelgrass	.0004195	.000024	17.35
Cost	-.003762	.000237	-15.91
Log-Likelihood	-4079.37		
Restricted Log-Likelihood	-4731.72		
Chi-Squared (Sig. Level)	1304.71 (.00000)		
$\rho^2=1-(L((\theta)/L(0))$	.138		

The nested logit model was structured so that upper branch of the nested model contains to the choice to take action or not, and the lower branch contains the choice of the specific program conditional on taking action. Thus, the willingness to pay for "taking action" *per se*, which we interpret as the symbolic component, is relevant only in the upper level of the decision tree in the nested model, where the individual chooses whether or not to "take action". This component of the choice is not relevant for the lower level of the decision tree, choosing between program A and program B given the decision to take action. This allows the nested model to isolate a source of symbolic value.

In the nested logit model, the probability that an individual chooses alternative k is

$$P_{kj} = P(k|j)P(j). \quad (2)$$

The choice probability for each of the three lowest level alternatives is conditional on the choice to take action or not. In this case, if the person chooses not to take action, then the probability of selecting the "No New Action" alternative is 1, since it is the sole alternative on that branch of the nested model. If they choose to take action, the probability of selecting "Program A" or "Program
B" is

$$P(k|j) = \frac{\exp(V_{k|j})}{\sum_{n|j}\exp(V_{n|j})} = \frac{\exp(\beta' x_{k|j})}{\sum_{n|j}\exp(\beta' x_{n|j})} = \frac{\exp(\beta' x_{k|j})}{\exp(J_j)}, \quad (3)$$

where k is one of the two alternatives, Program A or Program B; j is the choice to take action; n is the number of "action" alternatives (2); and Jj is the inclusive value for choice j, which represents the expected maximum utility from the choice of an alternative that sub-branch.

This is defined as

$$J_j = \log(\sum_{n|j}\exp(\beta' x_{n|j})). \quad (4)$$

The probability of choosing to take action or not is

$$P(j) = \frac{\exp(\alpha' Y_j + \tau_j J_j)}{\sum_m \exp(\alpha' Y_m + \tau_m J_m}, \quad (5)$$

where m represents each of the two branches.

The parameter τ is the inclusive value coefficient, which is related to the correlation between alternatives within a branch. A value of τ between 0 and 1 indicates that there is greater substitutability within, rather than across, groups of alternatives. This indicates that, even after accounting for the resources provided by the two programs, program A and program B are viewed by respondents as closer substitutes with each other than they are with the "no new action" alternative. Again, this is consistent with the notion that respondents view "taking action" as distinctly different from "no action" in ways that cannot be explained by the specific resources provided by those actions. We interpret this as a component of symbolic bias. In contrast, if τ is equal to 1, then all alternatives are equally substitutable, and the model becomes identical to the standard conditional logit model (McFadden 1981; Kling and Herriges 1995; Kling and Thomson 1996), which is consistent with the notion that respondents value the action alternatives only to the extent that they provide specific resource benefits.

As shown in Table 4, the results of both of these models indicate that there may be effects on choices unrelated to the described attributes. In the alternative-specific constant model, both constant terms are statistically significant, indicating that there is an effect on choices unrelated to the quantities of the individual attributes, but is instead related to the choices themselves. The positive and significant coefficient for "Action" indicates that people are more likely to choose an action rather than "No New Action," beyond what can be explained by of the action's "... specific levels of provision..." of natural resources. Similarly, the negative and significant coefficient for "Program B" indicates that, even if Program A and Program B produced the same results in terms of preservation of natural resources, respondents are more likely to choose Program A. The coefficient on Action indicates that there is an 87 percent probability that the average respondent would select one of the two action alternatives over no action, if the actions cost nothing and provided zero resource protection. Similarly, conditional on taking action, there is a 55 percent probability that the representative respondent would choose Program A over Program B if their costs and levels of resource protection were identical.

The constant term for taking action may be interpreted as representing a qualitative or symbolic dimension of respondents' preferences, while the coefficients on the natural resources represent the quantitative dimension that can be attributed to the stated levels of resource protection. Thus, if respondents exhibit a tendency to select a resource protection action, rather than "No New Action," beyond that which can be associated with the stated levels of resource protection and the cost, they may be expressing a symbolic willingness to pay to take action to protect the environment of the East End.

A possible alternative explanation for the constant term is a functional form specification error, resulting from the use of the linear function form in place of the "true", but unknown, nonlinear function. In order to test this hypothesis we also use a non-parametric specification, where the different levels of the attributes are represented by dummy variables, so that no functional form is imposed. The results for this non-parametric specification is shown in Table 5. As can be seen, all coefficients are of the correct sign and are statistically significant at the 1% level. Most importantly, the action dummy variable is large and statistically

Table 4 - Comparison of Model Results

	Coefficient	Value /Acre	Capitalized Value per Acre*	Estimated % Symbolic Value
Conditional Logit Model:				
Farmland	0.000511	$0.136	$116,286	
Undeveloped Land	0.000107	$0.028	$24,240	
Wetlands	0.000336	$0.089	$76,446	
Shellfish Areas	0.000233	$0.062	$ 53,086	
Eelgrass	0.000419	$0.111	$95,398	
Cost	-0.003765			
$\rho^2=1-(L(\beta)/L(0))$	.138			
Model with Alternative-Specific Constants:				
Action	1.2866			
Program B	-0.1799			
Farmland	.000300	$0.094	$80,081	31%
Undeveloped Land	.000057	$0.018	$15,193	37%
Wetlands	.000179	$0.056	$47,666	38%
Shellfish Areas	.000108	$0.034	$28,754	46%
Eelgrass	.000214	$0.067	$57,207	40%
Cost	-.003207			
ρ^2	.171			
Nested Logit Model:				
Program B	-.1586			
Farmland	.000300	$0.087	$74,562	36%
Undeveloped Land	.000056	$0.016	$14,024	42%
Wetlands	.000228	$0.066	$56,669	26%
Shellfish Areas	.000128	$0.037	$31,742	40%
Eelgrass	.000281	$0.082	$69,962	27%
Cost	-.003441			
τ	.3397			
ρ^2	.309			

* – Total capitalized value based on 73,423 households, a 7% discount rate and a 25 year planning horizon.

Table 5. Estimation Results for Non-Parametric Specification

Variable	Coefficient	Std Error	z=Coef/Std.Err.
Action	1.17	.086	13.56
Alternative B	-0.13	0.037	-3.60
Farmland-Medium	0.25	0.088	2.78
Farmland-High	.90	0.095	9.42
Undev. Land- Medium	.50	0.079	6.32
Undev. Land-High	.78	0.095	8.19
Wetlands-Medium	1.01	0.097	10.35
Wetlands-High	.90	0.111	8.09
Shellfish-Medium	0.27	0.089	2.98
Shellfish-High	0.44	0.090	4.87
Eelgrass-Medium	0.43	0.088	4.95
Eelgrass-High	0.68	0.087	7.82
Cost	-0.031	0.025	-12.40

significant in the non-parametric specification, indicating that the significant constant term is not an artifact of functional form specification.

The non-parametric version also allows to test for scope effects, whereby more inclusive commodities are preferred to less. With the exception of wetlands, all attributes pass the "scope test", since the high level of resource protection for each attribute has a larger coefficient estimate than the middle level, and statistically significant at the 10% level or better. The exceptional result regarding wetlands might be explained by concerns expressed in focus groups. The medium level of wetlands protection represents maintaining the 1995 level, while the high level of wetlands represents an increase in wetlands the 1995 level. During focus groups, some participants expressed the concern that these new wetlands might be located near their house, and that they could result in mosquito problems. This is consistent with the result that the medium level of wetlands protection has a large positive coefficient, while the high level of wetlands elicits a smaller coefficient, so that increasing wetlands beyond the 1995 level is viewed as less valuable than maintaining the current level of wetlands.

The symbolic effect of taking action is not surprising, given the level of concern among residents of the area for the environment of the Estuary. However, the preference of one program over another beyond the described effects is not expected, and could occur for a variety of reasons. For example, the effect could be related to the ordering of the two programs, their placement on the page, or could possibly indicate that respondents infer some preference from the labels (e.g., an "A" is better than a "B"). Note that this alternative-specific constant term, though statistically significant, is quantitatively small.

The use of a constant for the choice of action versus no action implies a fixed effect model, where the constant represents a mean "bias" towards action over no action, beyond that which can be explained by the described levels of resource protection and cost. An alternative approach to modeling is to use a random effects model, where the random components of preferences for the two action programs are correlated. This implies that an action/no action bias might exist, but that the bias is randomly distributed across choices. For example, some individuals might exhibit a bias towards taking action, while others might exhibit a bias against taking action. This might be the case, for example, if conservative respondents feel that the government is already excessively invasive, and should "stay out of our business".

The random effects model can be implemented using the nested logit approach, which captures the correlation of the random components of utility associated with the two action alternatives. Tests of the inclusive value parameter in the nested logit model indicate that there is greater substitutability between the two action alternatives than there is between either program and "no new action". The constant term for Program B is similar in magnitude to that estimated in the previous model.

Economic values for the conditional logit model were estimated based on Hanemann (1984), and are measured by the cost, C, that would make a person indifferent between the choice selected and the baseline, no action, which has zero cost. Thus, for the conditional logit model,

$$U(\mathrm{R}_{ij},M_i)=U(\mathrm{R}_{ik},M_i-C_k) \quad \text{for all } k\neq j, \tag{6}$$

where j represents the "No New Action" alternative, or the baseline levels of the resources, so that $C_j=0$; k is the option selected; and C_k is the maximum willingness to pay for option k. For the linear approximation of the utility function presented above, this can be solved for C_k as follows:

$$\beta\mathrm{R}_{ij}+\gamma(M_i)=\beta\mathrm{R}_{ik}+\gamma(M_i-C_k)\ \textit{and}\ C_k=-\frac{\beta}{\gamma}(\mathrm{R}_{ik}-\mathrm{R}_{ij}). \tag{7}$$

Thus, for the conditional logit model, the dollar value to the average respondent for a unit change in each of the natural resources is calculated as the ratio of the coefficient on the resource, β, to the coefficient on cost, γ.

The calculation of dollar values for the nested logit model must account for the nested structure and the inclusive value parameter, τ. The formula for the compensating variation associated with a change in one of the attributes of the choice is (Kling and Thomson 1996; Hanemann 1982):

$$CV=-\frac{1}{\gamma}\left\{\ln\left[\sum_{j=1}^{J}\left(\sum_{k=1}^{K_j}\exp(V_{jk}^{2}/\tau_j)\right)^{\tau_j}\right]-\ln\left[\sum_{j=1}^{J}\left(\sum_{k=1}^{K_j}\exp(V_{jk}^{1}/\tau_j)\right)^{\tau_j}\right]\right\} \tag{8}$$

where V is the utility function, the superscripts on V indicate whether the attributes are set at the new level or the old level, and γ is the coefficient on cost. The coefficients and estimated values for all three models are compared in Table 4.

The dollar values for the conditional logit model with alternative-specific constants and the nested logit model are similar, and are 26% to 46% smaller than those estimated from the base model. These values might be interpreted as the portion of respondents' WTP to take action which can be attributed to the described changes in natural resource levels. This is smaller than the estimated value in the base model, which includes the "symbolic" effect discussed above. The constant term for Program B in the nested logit model is slightly smaller than that estimated in the conditional logit model, although it is similar in magnitude.

The estimated dollar values and relative values, calculated as ratios between the coefficients on each pair of resources, were compared for each model using Friedman's test for more than two related samples (Neave and Worthington, 1988). Based on this test, the hypothesis of equality of the estimated dollar values for the three models is rejected. However, a comparison of the estimated dollar values for the nested logit model and the alternative-specific constants model using the Wilcoxon signed rank test does not reject the hypothesis of equality of values for these two models. In addition, the 95 percent confidence intervals for the values from these two models overlap.

The Friedman test does not reject the hypothesis that the relative values for natural resources are equal for all three models. Additionally, the ordinal priorities for all three models are the same, with farmland most important, followed by eelgrass, wetlands, shellfishing areas and undeveloped land. These results indicate that priorities and relative values are robust with respect to different model specifications, and are independent of symbolic effects, but that the estimated dollar values vary somewhat between the base model and the two alternative specifications. However, the estimated dollar values for the three models are close in magnitude. Therefore, that the proportion of value resulting from this "symbolic" effect, while significant, is not excessive.

The similarity of results from the nested logit and conditional logit with constants models, and the fact that the results are not statistically different, indicate that these biases are likely overwhelmingly in one direction—towards taking action rather than no action. Thus, both of these models appear to account for a "symbolic" aspect of values, and to separate that from estimated values for specific natural resource improvements. Note, however, the Nested Logit model provides considerable improvement in fit, as measured by the ρ^2 statistic.

IV. Summary and Conclusions

This paper describes an attempt to identify and correct for symbolic values during the data analysis process. The results of the study pass various statistical conditions for validity, including estimation of significant coefficients of the expected signs—positive for environmental attributes and negative for the cost of programs. Additionally, the results are quite robust to model specification, with equal priorities and relative values for different models, and dollar values that are relatively stable over different specifications, indicating that they are consistent with theory at a more rigorous level of validity. However, the results also indicate statistically significant symbolic effects.

Several steps were taken to test and correct for symbolic values. First, alternative specific dummy variables were used to identify tendencies to select action versus no action, and program A versus program B, beyond described the monetary costs and resource benefits of the programs. Monetary values of resources are then calculated net of these potentially symbolic effects.

Second, a nested logit model was developed, where responses are modeled based on a two level structure. The upper level of the decision tree contains the choice between action and no action. The lower level of the tree contains the chose between action A and action B, given that they chose to take action. This model implies that action A and action B are closer substitutes for each other than each is with no action. Once again, this could be consistent with symbolic bias to the extent that individuals view the two action alternatives as "similar", beyond what can be explained by the specific resource benefits described in the survey.

The results of the alternative-specific constants model and the nested logit model provide some evidence for the expression of symbolic values for the environment, beyond the values expressed for specific natural resources. Monetary values are similar for these two models, indicating that they both address symbolic effects equally well. Values for both the alternative-specific constants model and the nested logit model are 26% to 46% lower than those estimated with the standard conditional logit model. In comparison, relative resource values and resource priorities appear much better behaved, with relative resource values typically falling within 10% across the different models. This implies that symbolic effects do not have significant impacts on the relative values and resource priorities, so that all three models estimate these equally well.

Overall, these results suggest that the contingent choice method has considerable potential for estimating dollar values for natural resources, after accounting for symbolic effects. In addition, contingent choice offers the advantage of estimating relative values and priorities for natural resources, which are more robust to model selection in our case study. In some cases, informed policy choices do not require dollar measures of resource benefits. Rather for many social decisions, relative values of natural resources are sufficient. Thus, for policy analysis that requires prioritizing actions, or comparing relative values of actions, as well as for situations where there is concern about symbolic values, contingent choice methods may be a preferred method.

References

Addelman, Sidney. 1962a. "Orthogonal Main-Effect Plans for Asymmetrical Factorial Experiments," *Technometrics* 4 No. 1 (February): 21-46.

Addelman, Sidney. 1962b. "Symmetrical and Asymmetrical Fractional Factorial Plans," *Technometrics* 4 No. 1 (February): 47-57.

Addelman, Sidney, and O. Kempthorne. 1961. *Orthogonal Main Effect Plans*. United States Airforce, Office of Aerospace Research, Aeronautical Research Laboratory, ARL Technical Report 79.

Arrow, K.R., R. Solow, E. Leamer, P. Portney, R. Rander, and H. Schuman. "Report of the NOAA Panel on Contingent Valuation." Federal Register 58, No. 10 (1993): 4602-4614.

Carson, Richard T., Nicholas Flores and Norman Meade, 1997. "Contingent Valuation: Controversies and Evidence" UCSD Economics Discussion Paper 96-36, November 1996.

Desvouges, William H., F. Reed Johnson, Richard W. Dunford, Sara P. Hudson, K. Nicole Wilson and Kevin J. Boyle, 1993. "Measuring Natural Resource Damages with Contingent Valuation: Tests of Validity and Reliability" in Hausman *Contingent Valuation: A Critical Assessment* Elsevier Science Publishers.

Diamond, Peter A., Jerry A. Hausman, Gregory K. Leonard, and Mike A. Denning, 1993. Does Contingent Valuation Measure Preference? Experimental Evidence" in Hausman, Jerry A. *Contingent Valuation: A Critical Assessment* Amsterdam: North Holland.

East End Economic and Environmental Task Force of Long Island, New York. 1994. *Blueprint for Our Future: Creating Jobs, Preserving the Environment.* New York: Newmarket Press.

Economic Analysis, Inc. 1995. *The Peconic Bays System: A Perspective on Uses, Sectors, and Economic Impacts.* Report submitted to Peconic Estuary Program.

Greene, William H. 1993. *Econometric Analysis.* New York: Macmillan.

Greene, William H. 1995. *LIMDEP Version 7.0 User's Manual.* Bellport, NY: Econometric Software, Inc.

Hanemann, W. Michael. 1982. "Applied Welfare Analysis with Qualitative Response Models," California Agricultural Experiment Station, October.

Hanemann, W. Michael. 1984. "Welfare Evaluations in Contingent Valuation Experiments with Discrete Responses," *American Journal of Agricultural Economics* 66 (August): 332-341.

Hausman, Jerry, 1993. *Contingent Valuation: A Critical Assessment* Amsterdam: North Holland.

Kellert, Stephen. 1995 (forthcoming). "Environmental Values, Coastal Context, and a Sense of Place," in L. Anathea Brooks and Stacy VanDeveer, eds., *Saving the Seas: Science, Values, and International Governance.* New Haven: Yale University Press.

Kling, Catherine L., and Cynthia J. Thomson. 1996. "The Implications of Model Specification for Welfare Estimation in Nested Logit Models," *American Journal of Agricultural Economics* 78 No. 1 (February): 103-114.

Kling, Catherine L., and Joseph A. Herriges. 1995. "An Empirical Investigation of the Consistency of Nested Logit Models with Utility Maximization," *American Journal of Agricultural Economics* 77 No. 4 (November): 875-884.

Long Island Regional Planning Board. 1981. *Land Use: Quantification and Analysis of Land Use for Nassau and Suffolk County.*

Maddala, G.S. 1983. *Limited-Dependent and Qualitative Variables in Econometrics.* Cambridge: Cambridge University Press.

Mazzotta, Marisa J. and James J. Opaluch. 1995. "Decision Making When Choices are Complex: A Test of Heiner's Hypothesis," *Land Economics* 71, No. 4 (November): 500-515.

McFadden, Daniel. 1973. "Conditional Logit Analysis of Qualitative Choice Behavior," in P. Zarembka, ed., *Frontiers in Econometrics*. New York: Academic.

McFadden, Daniel. 1981. "Econometric Models of Probabilistic Choice," in C. F. Manski and D. McFadden, eds., *Structural Analysis of Discrete Data with Econometric Applications*. Cambridge: MIT Press.

Mitchell, Robert and Richard Carson, 1989. *Using Surveys to Value Public Goods: The Contingent Valuation Method* Washington, D.C..: Resources for the Future.

Ohio versus US Department of the Interior, 1989, 880 F.2d 432, 459 (D.C. Cir. 1989).

Sagoff, Mark. 1992b. "Settling America or The Concept of Place in Environmental Ethics," *Journal of Energy, Natural Resources and Environmental Law* 12 No. 2: 349-418.

Suffolk County Department of Health Services. 1992. *Brown Tide Comprehensive Assessment and Management Program.*

Suffolk County Department of Health Services. 1994. *Peconic Estuary Program Action Plan.* Suffolk County Department of Planning. 1991. *Estimated Peak Seasonal Population - 1990.* Unpublished spreadsheet.

U.S. Water Resources Council, 1983. *Economic and Environmental Principles and Guidelines for Water and Related Land Resources Implementation Studies* U.S. Government Printing Office, Washington D.C.

Regulatory Takings and the Diminution of Value: A Search for Empirical Evidence**

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Introduction

Although many zoning and regulatory takings cases have been seen in the courts over the past few years, the general public became much more concerned with regulatory takings after the well publicized case of *Lucas v. South Carolina Coastal Council*. The Court concluded that when legislation deprives an owner of all economically viable use of the property, compensation is required as long as the restriction was not originally part of the landowner's title. In *Pennsylvania Coal Co. v. Mahon*, Justice Homes gave the opinion that if a regulation "goes too far" it will be considered a compensable taking. This has been referred to as the diminution of value test: "how much value has been lost due to the regulation?" Following *Mahon*, the diminution of value test was used in a number of land mark takings cases including *Agins v. City of Tiburon, Deltona Corp. v. United States*, *Loveladies Harbor, Inc. v. United States* and *Florida Rock Industries, Inc. v. United States*.

Similarly, the House and the Senate have both put forth takings bills that set the level of diminution of value that determines when a taking has occurred. The bills, although different in values, were similar in their meaning: if an individuals property had a diminution of value equal to X percent due to a government action, then a compensable taking had occurred. Although neither of these bills were passed, there are very likely to come up again in the future.

It is apparent that the diminution of value is an important concept when discussing regulatory takings and compensation. But, how much has land devalued due to government environmental regulations and how might we measure that amount? Using data from Rockport and Corpus Christi, Texas, we are going to propose a methodology to test the affects of Section 404 of the Clean Water Act on land values. These values could be important if the above mentioned takings bills were to be passed. It would be useful for the courts to have a methodology to measure diminution of value as well as having a handle on the amounts that might be confronted. These values could also be used in a cost benefit analysis of Section 404 and the methodology used in other regulatory cost benefit analysis.

BACKGROUND AND THEORY

Parsons (1992) uses an hedonic property model to estimate property value changes in the Chesapeake Bay area resulting from regulations initiated by the Critical Areas Commission of Maryland. These regulations restrict land use on property abutting the Chesapeake Bay. All residential development within 1000 feet of the Bay's mean high tide line, designated as a critical area, is restricted. Some residential development is still permitted; however strict regulations incorporating landscaping, setback and impervious surface restrictions are now required. Also, new residential development is channeled into areas that are already developed and away from those less developed areas.

Using housing data in Anne Arundel County, Maryland, Parsons constructs a repeat sale analysis to estimate changes in property values before and after the regulation in the coastal area and several inland locations. He sets up an hedonic price function for the homes in the coastal area following Palmquist (1982) and Mendelsohn (1986) which includes attributes of the homes, dummy variables for two of the three critical area designations, a dummy for waterfront location, age of the house, and the distance the house is from the critical area if not located there. Each house in the data set was sold once before and once after the regulation. Price changes are in percentage terms and a real estate price index is incorporated. The data consisted of 441 transactions of which 31 percent were in the critical area. Several control areas were used each of which affected the percentage change in housing prices due to the effects of the regulation.

Parson's results how that housing prices increase as a result of the regulation of the critical areas. Homes in the critical areas had the largest effect, but homes as far away as 3 miles showed a 4 to 11 percent increase in price depending upon the control area used. Parsons contributes the price increased to the decreased supply of homes in the area. Apparently there are few substitutes for waterfront property since that is where the largest price increases were found. Moreover, homes even 3 miles away are affected by the decreased supply of homes in the critical area due to the new regulation. He also contributes some of the price increases to amenity values capitalized into the land and therefore into the housing prices from the perceived protection of coastal open space. Parsons concludes showing that others estimating changes in housing values for coastal areas with regulation changes offer results similar to his own.

Following Parsons' model, we formulate an hedonic model to estimate changes in land and housing prices in Rockport and Corpus Christi, Texas. Unlike many of the coastal regulations, Section 404 of the Clean Water Act strictly regulates whether land may or may not be developed. A land owner applies for a permit to develop a plot of land that is a wetland. The Army Corps of Engineers reviews the submission, delineates the land and then determines whether or not a permit will be granted. If not, the landowner is unable to develop a lot, and a significant decrease in value occurs. Lacking a permit, the lot is unlikely to be sold for very much, and maybe not at all. However, the value of surrounding property may increase due to the increasing scarcity of developable land.

We want to calculate the diminution of value caused by Section 404 of the CWA for those properties which have been denied a permit. We hypothesize that the price of non-permitted, undeveloped land will decrease and the price of permitted undeveloped land will increase after the incorporation of Section 404 permitting. The probable cause of the price increase is a decreased supply of developable land. The price effects in the housing market should be apparent as well. We should also see an increase in prices of homes in the area due to the decreased supply of developable land.

Since the introduction of 404 permitting, the law has gone through many changes. These changes had different effects including making the ability to obtain a permit easier and harder depending on which year we are studying. National permits were redefined which increased the maximum lot size that could obtain a national permit. During times in which it became easier to obtain a permit; it would be easier to develop undeveloped land. In that case, the supply of developable land would increase driving the price downward. On the other hand, if the permitting became more difficult, the supply of developable land would decrease driving the price upward. We are interested in measuring property value changes over time in response to these changes in 404 permitting.

DATA

The study area is part of the Texas coast which includes Corpus Christi and Rockport. These communities are approximately 35 miles apart and offer segregated land and housing markets. Individuals who choose to live in Corpus Christi do differ from those who choose to live in Rockport. Both of these areas are well established coastal communities that have been regulated heavily under Section 404 of the Clean Water Act. In both Corpus Christi and Rockport, there are a number of developments that were created before and after the 404 permitting began. We are interested in looking at homes and lots in these developments as well as homes and lots in developments that were not affected by the regulation.

Real estate sales data for homes and lots in Corpus Christi from 1972 to present were collected from the Corpus Christi Multiple Listing Service (MLS). Tax assessment values for homes and lots in Rockport from 1970 to 1985 were collected from the Rockport Tax Collection Office. Additional appraisal values and sales values for Rockport were collected from a local appraiser's office which included MLS sales values and appraisal values for sold properties from 1985 to present. Various additional sales values for Rockport and Corpus Christi were collected from a number of real estate agents in the area. Historical information on the real estate market in the area was obtained from a number of experienced real estate agents, developers and tax assessors.

Permit information was collected from the Corpus Christi and Galveston offices of the Army Corps of Engineers. Additional information was collected from offices in Dallas, Texas and Washington, D.C. We have information on which lots have permits and which do not and any restrictions, when applicable, on the properties. Amendments and changes to Section 404 of the Clean Water Act were taken from Federal Register documents obtained from the Army Corps of Engineers.

Hedonic Model

In formulating the hedonic model, we set up an hedonic price function for undeveloped land:

$$P^{UL} = P^{UL}(\mathbf{z}, W, V, D, M).$$

$\mathbf{z} = z(S,N,L)$ is a vector of structural characteristics S such as lot size, sewer and water, age; N is a vector of neighborhood characteristics such as quality of local schools, accessibility to stores and workplace, and local crime rate; and L is a vector of locational attributes such as distances to schools, stores, workplace, main roads, freeways, and waterfront. W is a dummy variable equal to 1 if the lot is located on the waterfront, V is a dummy variable equal to 1 if the lot has a view of the water, D is the distance the lot is from the water (Parsons and Wu, 1991) and M is a dummy variable equal to 1 if the lot has a permit. Any attribute's marginal implicit price is the first derivative of the equilibrium price function with respect to the attribute, or $\partial P/\partial z_i$, where z_i is any attribute in $P^{UL}(\mathbf{z},W,V,D,M)$. Therefore, if we are interested in measure the value of a permit, we can calculate $\partial P/\partial M$ which will give the marginal willingness to pay for a permit. Theory states that all else equal, two identical lots with permits should be the same price. It will be interesting to see if this is the case. Similarly, we can calculate the diminution of value for property without permits by:

$$\Delta P^{UL} = P^{UL}(z, W{=}1, V{=}1, D{=}d, M{=}1) - P^{UL}(a, W{=}1, V{=}1, D{=}d, M{=}0).$$

Finally, we will use the time series data to measure the property value changes through time given the set changes in the 404 permitting regulations. A dummy variable can be incorporated into the model for each of the time periods in which the regulations changed. If the coefficients on these dummies is significant, then effects of the regulations changes will be apparent. Additionally, we can segregate the data into different time periods for each of the regulatory changes. We can then compare the coefficient changes for each of the segregated data models.

On the consumer side, a buyer maximizes utility, $U(x,\mathbf{z},W,V,D,M,\alpha)$, by taking the equilibrium price function as exogenous. X is a numeraire good, α is a vector of socio-demographic characteristics and **z**, W, V, D, and M are the land characteristics contained in the hedonic price function. Utility is maximized subject to a budget constraint

$$Y = x + P(z_i,\ldots,z_n)$$

where Y is income. First order conditions require that

$$MRS_{zi,X} = \partial P/\partial z_i = U_{xi}/U_x.$$

This states that the consumer will purchase the attribute until the marginal implicit price of the attribute equals the marginal rate of substitution of the attribute for the numeraire good. The consumers bid function is the willingness to pay for the lot with **z** characteristics defined as

$$B(\mathbf{z}, U, Y; \alpha).$$

The bid function can be defined implicitly by

$$U(Y\text{-}B, z_1,\ldots,z_n;\alpha) = \mu.$$

If income changes, there is an equivalent change in the bid. Therefore in equilibrium, the marginal implicit price of the attribute will equal the consumers marginal bid or marginal willingness to pay for the attribute (Palmquist, 1991). If we can show that the demand for permits is identical to its marginal implicit price curve, we may be able to extract standard benefit measure such as compensating and equivalent surplus thereby evaluating the welfare affect of the lump sum change in wealth from not having a permit (Kriesel, et. al, 1993).

A similar hedonic price function will be constructed for developed land:

$$P^{DL} = P^{DL}(\mathbf{z},W,V,D).$$

In this case, the structural characteristics will include attributes specific to the home. These models for developed and undeveloped land will be estimated independently to isolate effects on developed and undeveloped land. Dummies for specific subdivisions may be necessary. In addition, we must take into account the economic changes that took place in this area. Corpus Christi and Rockport, like many other areas, went through a boom and bust period in the eighties which must be explicitly considered.

From a review of the literature, it appears that the double log functional form has been used frequently. Similarly linear models are estimated for a point of reference. More recently the linear Box-Cox form has become a common form for hedonic regressions (See Kriesel, et.al, 1993; Palmquist, 1984 and 1991; and Epple, D., 1987). It appears that theory predicts a non-linear model if costless repackaging is possible. Even though costless repackaging may be true for the housing market in the long run, we may not necessarily see this type of long run equilibrium in our model. It is therefore believed that the functional form for the hedonic model must be determined empirically (Palmquist, 1991). From reviewing the literature a number of functional forms have been proposed for hedonic equations. It appears that the double log and linear Box-Cox functional forms have performed quite well in a number of studies giving expected coefficient estimates with the expected signs (Parsons and Wu, 1991; and Kriesel, Randall and Lichtkoppler, 1993). We will consider both the double log and linear Box-Cox as important functional forms for this study. The permitting variable, M, may better enter the equation as a squared term given the importance of having a permit; however, that determination cannot be made at this time.

Since we have two segregated markets for land and homes, identification of demand functions should not be a problem. It is possible that we may have sample size problems with a time series model, but more examination of the data is necessary.

Unfortunately, aggregate census data will be used for consumer characteristics since individual data are not available

Implications and Conclusion

The objective of this study is to calculate how property values have changed over time due to Section 404 permitting regulations as well as to calculate the diminution of value caused by the regulation. Formulating an hedonic price function for developed and undeveloped land in conjunction with data collected from the Texas coast, we hope to prove our hypothesis. As regulations became stricter, more difficult to receive a permit, property values should increase due to the scarcity of developable land. Similarly, as regulations became less strict, easier to obtain a permit, property values should decrease due to the increased supply of developable land. Finally, land for which a permit was not obtainable should decrease significantly in value and we should be able to measure that decrease. There have been many theoretical models proposed, but no empirical models have been found on regulatory takings. Additionally, hedonic models have rarely been used to measure costs of regulations. The estimated values obtained can be used to measure the amount of compensation that would be necessary if 404 regulations in the Corpus Christi area were considered takings. These values will also allow the government to approximate the amount of money they would be dealing with if compensation had to be granted. The estimated values will also show the percentage of value lost due to the regulations which can be used to estimate when compensation would be required under different regimes outlined in the takings bills. When combined with estimates of public benefits for the regulations (e.g. WTP to preserve wetlands), the values can be used in a benefit cost analysis of the 404 regulation. Finally, we may be able to use the hedonic values as approximations of Hicksian welfare measure or attempt to derive Hicksian measure from the hedonic model. The bottom line is that it may not be as costly to compensate individuals for their losses as previously believed.

References

Blume, Lawrence, and Daniel L. Rubinfeld. "Compensation for Takings: An Economic Analysis." The economics of urban property rights. Research in Law and Economics series, vol. 10, Greenwich, Conn. Ed. Austin- J. ed. Jaffe. and London: JAI Press, **1987**. 53-103.

Brown, G.M. and H.O. Pollakowski. "Economic Valuation of Shoreline." The Review of Economics and Statistics 59 (2), **1977**: 272-278.

Diamond, D.B. and B.A. Smith. "Simultaneity in the Market for Housing Characteristics."
Journal of Urban Economics 17 (2), **1985**: 280-292.

Edwards, S. and G. Anderson. "Land Use Conflicts in the Coastal Zone: An Approach for the Analysis of the Opportunity Costs of Protecting Coastal Resources." Journal of Northeastern Agricultural Economics April **1984**: 73-81.

Epple, D. "Hedonic Prices and Implicit Markets: Estimating Demand and Supply Functions for Differentiated Products." Journal of Political Economy 95 (1), **1987**: 59-79.

Fischel, William A. Regulatory takings: Law, economics, and politics. Cambridge and London: Harvard University Press, **1995**.

Grieson, Ronald E. and James R. White. "The Effects of Zoning on Structure and Land Markets." Journal of Urban Economics 10 **1981**: 271-285.

Kriesel, Warren, Alan Randall and Frank Lichtkoppler. "Estimating the Benefits of Shore Erosion Protection in Ohio's Lake Erie Housing Market." Water Resources Research 29 (4), **1993**: 795-801.

Mendelsohn, T.A. "Review of Identification of Hedonic Supply and Demand Functions." Growth and Change 18 (1), **1987**: 82-92.

Miceli, Thomas J. "Compensation for the Taking of Land under Eminent Domain." Journal of Institutional and Theoretical Economics 147 (2), **1991**: 354-63.

Miceli, Thomas J., and Kathleen Segerson. "Regulatory Takings: When Should Compensation Be Paid?" Journal of Legal Studies 23 (2), **1994**: 749-76.

Michelman, Frank I. "Property, Utility, and Fairness: Comments on the Ethical Foundations of "Just Compensation" Law." Harvard Law Review 80 (2), **1967**: 1165-1258.

Milton, W.J., J. Gressel, and D. Mulkey. "Hedonic Amenity Valuation and Functional Form Specification." Land Economics 60 (Nov), **1984**: 378-387.

Munch, Patricia. "An Economic Analysis of Eminent Domain." Journal of Political Economy 84 (3), **1976**: 473-497.

Nelson, Robert H. "Federal Zoning: The New Era in Environmental Policy." Land Rights, **1990**. 295-317.

Palmquist, R.B. "Alternative Techniques for Developing Real Estate Price Indexes." Review of Economics and Statistics 62 (2), **1980**: 442-448.

Palmquist, R.B. "Estimating the Demand for the Characteristics of Housing." Review of Economics and Statistics 65 (5), **1984**: 394-404.

Palmquist, Raymond B. "Welfare Measurement for Environmental Improvements Using the Hedonic Model: The Case of Nonparametric Marginal Prices." Journal of Environmental Economics and Management 15 **1988**: 197-312.

Palmquist, Raymond B. "Hedonic Methods." Measuring the Demand for Environmental Quality. Ed. J.B. Braden and C.D. Kolstad. New York: North-Holland, **1991**. 77-120.

Parsons, George R. "Hedonic Prices and Public Goods: An Argument for Weighting Locational Attributes in Hedonic Regressions by Lot Size." Journal of Urban Economics 27 **1990**: 308-321.

Parsons, George R., and Yangru Wu. "The Opportunity Cost of Coastal Land Use Controls: An Empirical Analysis." Land Economics 67 (3), **1991**: 308-316.

Parsons, George R. "The Effect of Coastal Land Use Restrictions on Housing Prices: A Repeat Sale Analysis." Journal of Environmental Economics and Management 22 **1992**: 25-37.

Polinsky, A.M. and Steven Shavell. "Amenities and Property Values in a Model of an Urban Area." Journal of Public Economics 5 **1976**: 119-129.

Rosen, Sherwin. "Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition." Journal of Political Economy 82 **1974**: 34-55.

Runge, C. Ford. "Economic Implications of Wider Compensation for "Takings" or, What if Agricultural Policies Ruled the World?" Vermont Law Review 17 **1993**: 723-739.

Shabman, L. and M.K. Bertelson. "The Use of Development Value Estimates for Coastal Wetland Permit Decisions." Land Economics 55 (May), **1979**: 213-222.

Shapiro, Perry, and William A. Fischel. "Takings, Insurance and Michelman: Comments on Economic Interpretations of 'Just Compensation' Law." Trans. : U CA, Santa Barbara; Dartmouth College, 1987.

Shapiro, Perry, and William A. Fischel. "A Constitutional Choice Model of Takings." Trans. : U CA, Santa Barbara; Dartmouth College, 1987.

Unknown. "The Takings Defense." Law of Wetlands Regulation. Chapter 10, p. 1-15.

Unknown. "Government Operations, House Passes Takings Compensation Bill, Clears HR 9 Regulatory Relief Package." Trans. . Washington, D.C.: The Bureau of National Affairs, Inc., 1995.

Optimal Pollution Taxes in a Second-Best World

by

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Abstract

This paper models the social cost of an environmental policy in an economy with pre-existing tax distortions. The model is able to account for welfare impacts not included in previous work. With this model, the effect of relaxing the assumption that utility is separable between the environmental amenities and commodities is also analyzed. Results show that, without separability between the externality and market goods, the pigouvian tax is not optimal in a first-best world. Results also show that, without separability and with pre-existing tax distortions, the optimal pollution tax will often be below but can be above the pigouvian tax.

Introduction

Pigou showed that, in the existence of a consumption externality, the optimal tax or subsidy must equal to the marginal external cost or benefit. Recently the literature has begun to reassess the pigouvian tax/subsidy because, in an economy with pre-existing taxes, environmental taxes were seen to provide a second benefit or double dividend. This second benefit results from using the environmental tax revenues to reduce the pre-existing tax(es). The welfare gain from the decrease in the pre-existing tax(es) has been referred to as the "revenue recycling effect".

Unlike the first-best world of Pigou, environmental taxes in a second-best world also impose a "tax interaction" effect (Goulder). With some limiting market assumptions, Bovenberg and de Mooij show that, with the existence of a labor tax, the tax interaction effect is enough to drive the optimal environmental tax below the pigouvian tax. Parry results are consistent with these while using a slightly less restrictive model. However, these earlier models do not included all welfare impacts of the environmental tax.

This paper offers a model for assessing the full welfare impact of environmental policies in an economy with pre-existing tax distortions. Analysis shows that both benefits and costs have been excluded. Consequently, conclusions follow those of Parry and Bovenberg and de Mooij.

This paper also considers the welfare impact of relaxing the assumption that utility is separable between the environmental amenity and commodities. Results show that, without separability, the pigouvian tax is not optimal in a first-best world; the optimal pollution tax is even lower in a second-best world when the externality directly affects consumption.

Recent work and work here provide results that are consistent with the optimal commodity tax literature. Tax efficiency is assumed to exist prior to recognition of the externality. Obviously, tax efficiency does not exist in the real world and, therefore, could serve as a justification for changing the tax mix. However, tax efficiency is a separable issue (Bohm). By assuming tax efficiency, the model can assess the tradeoff between the benefits of the environmental tax relative to the social cost of moving from tax efficiency.[1]

The paper begins by reviewing the behavioral and market assumptions of previous work. Consistent with assumptions of previous work, a model of the welfare impacts of an environmental tax is derived. The model is first used to determine the optimal environmental tax in an economy with pre-existing tax distortions. The model is then used to determine the optimal environmental tax in an economy where there are no pre-existing distortions but there no separability between the environmental externality and consumption of market goods. Finally, conditions of pre-existing taxes and no separability are both applied.

Behavioral and Market Assumptions and Modeling Framework

As discussed in the introduction, the behavioral assumptions employed here have been employed in earlier related work (Parry; Bovenberg and de Mooij; Goulder, Parry, and Burtraw). To begin with, tax efficiency is assumed to exist before a pollution tax is introduced. Output from the polluting sector is assumed to be an average substitute for leisure.[2] Consumer preferences are

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[1]With tax efficiency, the aggregate tax burden or deadweight loss, associated with the a given level of tax collections, has been minimized (Sandmo, 1976). It should be noted that distributional considerations, considered in some work on optimal tax systems, are not considered here to simplify the analysis but with no loss in generalities.

[2]Parry notes this as a restrictive assumption of Bovenberg and de Mooij in his introduction yet is forced to apply this assumption in equation 11. However, in order for a labor tax to be efficient, the income elasticity of demand for all goods must be unity (e.g. homothetic preferences) and consumer preferences must be separable between consumption and labor/leisure (Atkinson and Stiglitz). Alternatively, a labor tax is efficient if labor supply is completely inelastic.

assumed to be homothetic. Production technologies are assumed to be linear and the only factor of production is assumed to be labor. Labor supply is assumed to be a function of the real wage (or the utility from consumption) and the dis-utility associated with working. The labor-leisure decision views individuals as trading off the dis-utility of labor with the utility gained from the consumption of goods purchased with the labor income. Thus consumers' preferences are separable between labor and all goods so that changes in labor do not affect the marginal rate of substitution between goods (Sandmo, 1976, and Atkinson and Stiglitz). Real incomes and government expenditures are held constant so that leisure and consumption decisions are measured with compensated demands. Cross-substitution responses are assumed to remain constant.

The government is assumed to maintained a balanced budget through lump-sum transfers. There are no government expenditures. Capital markets are not affected because all tax changes are viewed as permanent and the polluting industry produces only a consumption good. Markets are viewed as competitive and the economy is assumed to be void of economic rents. Marginal costs are assumed to be constant in each market.

As first shown by Harberger, the effect of taxes on social welfare, W, or the negative of the dead weight loss (DWL) associated with taxes is:[3]

$$W = \frac{1}{2}\left(t_X t_L \frac{\partial X}{\partial t_L} - t_X t_L \frac{\partial l}{\partial t_X} + t_X^2 \frac{\partial X}{\partial t_X} - t_L^2 \frac{\partial l}{\partial t_L} \right) \qquad (1)$$

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[3]As mentioned earlier, real incomes are held constant. Thus the DWL triangles are those of the compensated demands. Also note that leisure demand is substituted for labor supply. Since $\partial L/\partial t_L = -\partial l/\partial t_L$, the signs in equations 1 are adjusted from those in Harberger. Furthermore, supply and demand curves are assumed to be linear over the relevant range.

where W, the negative of the aggregate tax burden, is measured in terms of real purchasing power. Tax efficiency implies W is maximized at the level of tax revenue collected. Labor, leisure, a consumptive good, a tax on L, and a tax on X are represented by L, l, X, t_L, and t_X.

Holding tax collections constant, the change in welfare associated with balanced-budget changes in t_L and t_X is:

$$dW = \frac{1}{2}\left(t_X\frac{\partial X}{\partial t_L}dt_L - t_X\frac{\partial l}{\partial t_X}dt_L - t_L\frac{\partial l}{\partial t_X}dt_X + t_L\frac{\partial X}{\partial t_L}dt_X \right) + t_X\frac{\partial X}{\partial t_X}dt_X - t_L\frac{\partial l}{\partial t_L}dt_L. \quad (2)$$

The welfare change, dW depends on, among other things, the relative sizes of t_L, t_X, dt_L, and dt_X.

The integratability condition:

$$\frac{\partial X}{\partial t_L} = -\frac{\partial l}{\partial t_X} \quad (3)$$

allows equation 2 to be simplified to:

$$dW = t_X\frac{\partial X}{\partial t_L}dt_L - t_L\frac{\partial l}{\partial t_X}dt_X + t_X\frac{\partial X}{\partial t_X}dt_X - t_L\frac{\partial l}{\partial t_L}dt_L. \quad (4)$$

Based on the relationship provided by Harberger, the social cost of a tax on labor and a tax on commodity X are easy to see. With increases (decreases) in t_X, $t_X(\partial X/\partial t_X)dt_X$ and $t_L(\partial l/\partial t_X)dt_X$ are negative (positive) indicating a decrease (increase) in welfare. Likewise, with increases (decreases) in t_L, $t_X(\partial X/\partial t_L)dt_L$ and $t_L(\partial l/\partial t_L)dt_L$ are negative (positive) indicating a decrease (increase) in welfare. With tax efficiency, a marginal balanced-budget change in both t_X and t_L

will impose no welfare cost (e.g. $dW=0$) -- by definition. As one moves away from tax-efficiency with further substitutions of one tax for the other, dW becomes negative.

Equation 4 represents a comprehensive measure of the tax burden of two taxes. Previous work has explicitly excluded $t_X(\partial X/\partial t_L)dt_L$ arguing that the effect is relatively small (Parry). However small, inclusion of this term allows the analysis here to be more comprehensive.

The Trade-off Between Tax Costs and Environmental Benefits

If E_X is defined as the nonmarket cost associated with consuming one unit of X, then Pigou suggested that social welfare is maximized when $p_X(\partial X/\partial t_X)dt_X = E_X(\partial X/\partial t_X)dt_X$ or $p_X = E_X$. After increasing the tax on X by p_X, further increases will be optimal when $dW > E_X(\partial X/\partial t_X)dt_X$ or:

$$0 < (t_X - p_X)\frac{\partial X}{\partial t_X}dt_X - t_L\frac{\partial l}{\partial t_X}dt_X + t_X\frac{\partial X}{\partial t_L}dt_L - t_L\frac{\partial l}{\partial t_L}dt_L. \tag{5}$$

For simplicity and to be consistent with Parry and Bovenberg and de Mooij, t_X is assume to equal zero before implementing the externality tax. Thus the first term on the right-hand-side of equation 5 will equal zero. Equation 5 holds if the sum last three terms is greater than zero. With an increase in t_X and cut in t_L we know that one term is negative and the others positive. But it is not clear whether or not equation 5 holds. A look at the dynamics provides insight.

In order for equation 5 to hold, or $dW > E_X(\partial X/\partial t_X)dt_X$, then:

$$d^2W > dt_X\frac{\partial X}{\partial t_X}dt_X + t_X\frac{\partial X}{\partial t_X}dt_X^2 \tag{6}$$

as t_X is being increased to p_X. That is, the rate of increase in the total social costs of the taxes, d^2W, must be less than the rate of increase in the direct social cost of the tax on X, at least for some range of t_X.

By differentiating equation 2 we know that:

$$d^2W = dt_X \frac{\partial X}{\partial t_L} dt_L + t_X \frac{\partial X}{\partial t_L} dt_L^2 - dt_L \frac{\partial l}{\partial t_X} dt_X - t_L \frac{\partial l}{\partial t_X} dt_X^2 + dt_X \frac{\partial X}{\partial t_X} dt_X \qquad (7)$$
$$+ t_X \frac{\partial X}{\partial t_X} dt_X^2 - dt_L \frac{\partial l}{\partial t_L} dt_L - t_L \frac{\partial l}{\partial t_L} dt_L^2$$

where the first four terms quantify the dynamics of the cross-substitution terms and the last four terms quantify the dynamics of the direct affects of the tax changes.[4] To simplify, marginal changes in t_L, dt_L, are held constant, which is reasonable while the operating point on the X-Laffer curve is not very flat. With $dt_L^2 = 0$ and equation 7 substituted into equation 6, we have:

$$dt_L \frac{\partial l}{\partial t_L} dt_L + t_L \frac{\partial l}{\partial t_X} dt_X^2 < dt_X \frac{\partial X}{\partial t_L} dt_L - dt_L \frac{\partial l}{\partial t_X} dt_X. \qquad (8)$$

Since $dt_L(\partial l/\partial t_L)dt_L < -dt_L(\partial l/\partial t_X)dt_X$ (e.g. the leisure increase due to dt_X, $(\partial l/\partial t_X)dt_X$, is of greater magnitude than the decrease due to dt_L, $(\partial l/\partial t_L)dt_L$[5]), equation 8 will hold when $dt_X(\partial X/\partial t_L)dt_L >$

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[4]With an increase in t_X and decrease in t_L, the signs of terms are:

$$d^2W = dt_X \frac{\partial X}{\partial t_L} dt_L + t_X \frac{\partial X}{\partial t_L} dt_L^2 - dt_L \frac{\partial l}{\partial t_X} dt_X - t_L \frac{\partial l}{\partial t_X} dt_X^2 + dt_X \frac{\partial X}{\partial t_X} dt_X + t_X \frac{\partial X}{\partial t_X} dt_X^2 - dt_L \frac{\partial l}{\partial t_L} dt_L - t_L \frac{\partial l}{\partial t_L} dt_L^2$$

[5]Parry refers to this as the welfare loss due to the leisure market distortion.

$t_L(\partial l/\partial t_X)dt_X^2$ or $-dt_L/t_L > dt_X^2/dt_X$. And $-dt_L/t_L > dt_X^2/dt_X$ is more likely to hold:

A) the smaller the labor tax,

B) the steeper the X-Laffer curve at the operating points thus the smaller dt_X^2, and

C) the larger dt_X is relative to dt_L or the greater the share X composes of total commodity demand.

This result suggests that society can gain enough from the reduction in the deadweight loss of the income tax and the increase in environmental benefits so that the optimal environmental tax can exceed the pigouvian tax. However, there is an additional welfare impact not yet included.

Optimal Pollution Tax Recognizing the Environmental Impact of a Labor Tax

This section recognizes the effect the labor tax has on the externality. With externality costs given as $E_X X$, the effect of a balanced-budget change in taxes on externality costs is $E_X(\partial X/\partial t_X)dt_X + E_X(\partial X/\partial t_L)dt_L$. The total marginal welfare impact from a balanced-budget change in taxes, equation 4, minus the change these externality cost is:

$$t_X\frac{\partial X}{\partial t_L}dt_L - t_L\frac{\partial l}{\partial t_X}dt_X + t_X\frac{\partial X}{\partial t_X}dt_X - t_L\frac{\partial l}{\partial t_L}dt_L - E_X\frac{\partial X}{\partial t_L}dt_L - E_X\frac{\partial X}{\partial t_X}dt_X \qquad (9)$$

The pigouvian tax is suboptimal if the total marginal welfare change is positive when $t_X = E_X$, or:

$$0 < -t_L\frac{\partial l}{\partial t_X}dt_X - t_L\frac{\partial l}{\partial t_L}dt_L. \qquad (10)$$

This condition will not hold since $(\partial l/\partial t_X)dt_X > -(\partial l/\partial t_L)dt_L$ as discussed above. Under these conditions, the optimal pollution tax must be less than the pigouvian tax. This condition differs from the condition given in equation 5 by recognizing the added cost of the externality that results as the labor tax is reduced and consumption of X increases.

Equation 9 helps to explain the net loss in welfare associated with a marginal increase in the pigouvian tax. It is clear that a marginal increase in the tax on X provides a direct environmental gain equal to the tax's direct welfare cost ($E_X(\partial X/\partial t_X)dt_X = t_X(\partial X/\partial t_X)dt_X$). However, what is revealed here is that the labor tax cut provides a gain in welfare $t_X(\partial X/\partial t_L)dt_L$ that is offset by an associated increase in externality, $E_X(\partial X/\partial t_L)dt_L$. What remains are the direct gain in welfare associated with the reduction in the labor tax, $-t_L(\partial l/\partial t_L)dt_L$, and the smaller indirect cost in the labor market associated with the increase in t_X, $-t_L(\partial l/\partial t_X)dt_X$.

It is interesting to note that while Parry and Bovenberg and deMooij accounted for neither $t_X(\partial X/\partial t_L)dt_L$ nor $E_X(\partial X/\partial t_L)dt_L$, the net effect provided non-distortionary results. Thus, although the investigation here is more comprehensive, it leads to the same conclusion: that a pigouvian tax exceeds the optimal environmental tax.

Optimal Pollution Tax With Behavioral Responses to an Externality

In reality, utility is often not separable in environmental quality although this is commonly assumed in theoretical and applied research. In these cases, the externality affects what Randall and Stoll (1983) identified as use values (as opposed to nonuse values). When an externality affects use values, there is a market response to a change in the externality. For example, changes in water quality, air quality, and wildlife habitat affect expenditures for market goods associated with water-base recreation, health and visual amenities, and wildlife hunting and viewing activities, respectively. In these cases, changes in an externality affect either labor supply or commodity demands which, in turn, affect tax collections.

In the discussion below, utility is not assumed to be separable in environmental quality and the subsequent tax efficiency relationships are formalized. The impacts of this assumption are demonstrated, first, in an economy with only a single commodity tax, second, in an economy with a single labor tax, and, finally, in an economy with both the commodity and labor taxes. In this last case, the optimal pollution tax is assessed and results compared to the above findings.

The Optimal Pollution Tax in a First-Best World

This section shows that when utility is not separable in environmental quality the pigouvian tax is not optimal. This is because the demand for the polluting commodity is indirectly affected by the externality's impact on the second commodity. The externality could have been assumed to affected labor supply (or leisure demand) instead of a consumption good but conclusions would not differ.

Commodity X is again assumed to generate the externality. For each unit of X consumed, the marginal social cost is greater (less) than the marginal private cost by E_X for a negative (positive) externality where E_X is, again, the difference between the marginal social and marginal private costs of consuming X. The total cost of a negative externality in the X-market, $E_X X$, is the shaded area in figure 1.

The externality directly affects the utility of consuming good z.[6] Thus its cost can be measured in the market for z. The cost of the externality is equivalent to a tax on the consumption of z. The social cost of the externality in the z-market is equivalent to the total (gross) social cost of a tax on z (figure 2). Because the welfare subsidy in the X market must equal the welfare cost in the z market, we know that:

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[6]This is assumed to be the only welfare cost of the externality.

$$E_X X = \int_{MC_z}^{MC_z+E_z} D_z dz = E_z(z+\frac{1}{2}\frac{\partial z}{\partial E_z}E_z) \qquad (11)$$

where MC_z is the marginal private cost of z, assumed constant, and D_z is the demand for z, assumed to be linear over the relevant range. Furthermore E_z is the decrease in the utility of consuming (or using) a unit of z due to the externality and is a function of the total X consumed.

The change in the social cost of an externality associated with a marginal change a pollution tax can be measured in both the X and z markets:

$$E_X\frac{\partial X}{\partial t_X}dt_X+X\frac{\partial E_X}{\partial z}\frac{\partial z}{\partial t_X}dt_X = (z+E_z\frac{\partial z}{\partial E_z})\frac{\partial E_z}{\partial X}\frac{\partial X}{\partial t_X}dt_X+E_z\frac{\partial z}{\partial t_X}dt_X. \qquad (12)$$

The first terms on both sides of the equation 12 represent the change in the direct marginal social cost of the externality associated with a change in a tax on X. The first term on the left-hand-side, $E_X(\partial X/\partial t_X)dt_X$, is the traditional measure employed and is represented as area A_X in figure 3. This same welfare change is represented in the z-market by area A_z in figure 4.

The second terms represent the indirect or market interaction effect. This might be thought of as a 'price effect' on the total social cost of consuming X. The indirect effect reflects the shift in the demand for z in response to the change in the after-tax price of X. Since we've assumed X and z are substitutes, an increase in t_X increases the demand for z and decreases the benefit of the environmental tax by $X(\partial E_X/\partial z)(\partial z/\partial t_X)dt_X$ which is represented by the area B_X in figure 5. In the z-market, the increased demand for z increases welfare loss from the externality by $E_z(\partial z/\partial t_X)dt_X$ or area the B_z in figure 6. If the externality affected only nonuse values, then there would be no market interaction effect.

Under these conditions, the total effect on welfare of an environmental tax, based on equations 1 and 11, is:

$$W = \frac{1}{2}\left(t_X^2\frac{\partial X}{\partial t_X}\right) - E_z(z + \frac{1}{2}\frac{\partial z}{\partial E_z}E_z). \quad \textbf{(13)}$$

The change in welfare associated with a change in t_X is:

$$dW = t_X\frac{\partial X}{\partial t_X}dt_X - E_X\frac{\partial X}{\partial t_X}dt_X - E_z\frac{\partial z}{\partial t_X}dt_X \quad \textbf{(14)}$$

and is maximized when dW equals zero. Therefore, when an externality has a market impact, the optimal pollution tax is less (greater) than the pigouvian tax when X and z are substitutes (complements) because $\partial z/\partial t_X$ will be positive (negative). However if X and z were complements, the preliminary tax efficiency conditions would be violated.

Should the externality directly and adversely affect the decision to work (such as poor air quality concentrated in the factory or business area), then the optimal pollution tax is greater than the pigouvian tax. This is because the direct market impact of the externality is on labor and not on demand for a commodity. In such this case, substitute $E_l(\partial l/\partial t_X)dt_X$ for $E_z(\partial z/\partial t_X)dt_X$ in equation 14 and note that $-E_l(\partial l/\partial t_X)dt_X$ has, converse to $-E_z(\partial z/\partial t_X)dt_X$, a positive welfare impact. Thus the optimal pollution tax will exceed the pigouvian tax.

The Marginal Cost of the Labor Tax in a First-Best World

The change in the social cost of the externality associated with a marginal change in the labor tax is derived by differentiating equation 11:

$$E_X\frac{\partial X}{\partial t_L}dt_L+X\frac{\partial E_X}{\partial z}\frac{\partial z}{\partial t_L}dt_L = (z+E_z\frac{\partial z}{\partial E_z})\frac{\partial E_z}{\partial X}\frac{\partial X}{\partial t_L}dt_L+E_z\frac{\partial z}{\partial t_L}dt_L. \tag{15}$$

where the left-hand terms are measures of welfare changes in the X-market and the right-hand terms are the same measures in the z-market. The first and second right-hand terms equal the first and second left-hand terms, respectively, as in equation 12. Here, however, both $\partial X/\partial t_L$ and $\partial z/\partial t_L$ are negative so that an increase (decrease) in the labor tax will decrease (increase) the social cost of the externality.

The total welfare impact of the labor tax is:

$$W = -\frac{1}{2}\left(t_L^2\frac{\partial l}{\partial t_L}\right)-E_z(z+\frac{1}{2}\frac{\partial z}{\partial E_z}E_z). \tag{16}$$

The social cost of a change in the labor tax is:

$$dW = -t_L\frac{\partial l}{\partial t_L}dt_L-E_X\frac{\partial X}{\partial t_L}dt_L-E_z\frac{\partial z}{\partial t_L}dt_L. \tag{17}$$

Thus marginal social cost of an increase in the labor tax has a lower (higher) social cost when there is a negative (positive) externality in the economy and the externality affects use values.

Should the externality, again, directly and adversely affect the decision to work, then the cost of the pollution tax is reduced. To see this, substitute $E_l(\partial l/\partial t_L)dt_L$ for $E_z(\partial z/\partial t_L)dt_L$ in equation 17 and note that $-E_l(\partial l/\partial t_L)dt_L$ has, converse to $-E_z(\partial z/\partial t_X)dt_X$, a negative welfare impact.

The Pollution Tax in a Second-Best World

With both t_X, t_L, and an externality on the consumption of z generated by the consumption of X, the total welfare function is:

$$W = \frac{1}{2}\left(t_X t_L \frac{\partial X}{\partial t_L} - t_X t_L \frac{\partial l}{\partial t_X} + t_X^2 \frac{\partial X}{\partial t_X} - t_L^2 \frac{\partial l}{\partial t_L}\right) - E_z(z + \frac{1}{2}\frac{\partial z}{\partial E_z}E_z). \qquad \textbf{(18)}$$

With an increase in t_X offset by a decrease in t_L, the total marginal welfare impact is:

$$dW = t_X \frac{\partial X}{\partial t_L}dt_L - t_L \frac{\partial l}{\partial t_X}dt_X + t_X \frac{\partial X}{\partial t_X}dt_X - t_L \frac{\partial l}{\partial t_L}dt_L - E_X \frac{\partial X}{\partial t_X}dt_X - E_X \frac{\partial X}{\partial t_L}dt_L - E_z \frac{\partial z}{\partial t_X}dt_X - E_z \frac{\partial z}{\partial t_L}dt_L \qquad \textbf{(19)}$$

which follows from equations 4, 14, and 17. The first six terms are the same as in equation 9. Since a marginal increase in t_X and decrease in t_z leaves both $-E_z(\partial z/\partial t_X)dt_X$ and $-E_z(\partial z/\partial t_L)dt_L$ negative, the optimal pollution tax will be even lower than indicated in equation 10 (where there was no behavioral response to the externality).

Should the externality directly and adversely affect the decision to work, then the welfare cost of the pollution tax is reduced. This reduction in welfare is significant enough to make the pigouvian tax suboptimal since X and z are average goods. Specifically, as average goods, X=z (e.g. the quantities of X and z consumed are equal) and E_X=E_z (so that equation 11 holds). A change in the labor tax will have the same effect on X as z so that $(\partial X/\partial t_L)dt_L$ equals $(\partial z/\partial t_L)dt_L$ and thus $E_X(\partial X/\partial t_L)dt_L$=$E_z(\partial z/\partial t_L)dt_L$. This means that the welfare impact of equation 19 can be reduced to:

$$dW = t_X \frac{\partial X}{\partial t_L} dt_L - t_L \frac{\partial l}{\partial t_X} dt_X - t_L \frac{\partial l}{\partial t_L} dt_L - E_z \frac{\partial z}{\partial t_X} dt_X \qquad (20)$$

which has one addition term over equation 5 (given t_X equals p_X). This term, $-E_z(\partial z/\partial t_L)dt_L$, represents a positive change in welfare. Therefore the optimal pollution tax will be greater than the pigouvian tax across a wider range of conditions than those relevant to equation 5.

Summary

A model is developed to determine optimal environmental taxes. The model, first, is used to reexamine findings of previous work but includes a more comprehensive assessment of costs and benefits. Then, after dropping the assumption that utility is separable in environmental quality, the model is used to determine optimal pollution taxes in both a first-best and a second-best world. Results also show the conditions where the optimal pollution tax exceeds the pigouvian tax.

The model recognizes the increase environmental cost possible with a decrease in the labor tax. Although relevant in previous analyses, this impact has been ignored. The model also accounts for the welfare gains in the goods market associated with a decrease in the labor tax. While pointed out as a distortion in previous work, this welfare impact was not accounted for. Together, the additional externality cost and tax benefit are shown to be offsetting. Therefore, with a more comprehensive assessment of costs and benefits, the model shows that the optimal balanced-budget change in pollution tax in a second-best world is consistent with findings in previous research.

In the next stage of this analysis, the model is used to show that, when the externality has market impacts (specifically, the externality affects the demand for a consumed good) in a first-best or pigouvian world, the optimal pollution tax is less than the pigouvian tax. Specifically, a marginal increase in the pollution tax increases the demand for all other goods which, in turn, increases the demand(s) for the good(s) affected by the externality; the demand increase(s) increase the social cost of producing the polluting good.

In a second-best world, where an externality has market impacts, the optimal pollution tax is even further below the pigouvian tax. This is because both the market impact of the externality and secondary impacts of the environmental tax reduce the net marginal benefits.

References

Atkinson, A.B. "Optimal Taxation and the Direct Versus Indirect Tax Controversy," Canadian J. of Economics, Nov. 1977, 10, 590-606.

Atkinson, A.B. and J.E. Stiglitz. "The Structure of Indirect Taxation and Economic Efficiency," J. Of Public Economics, April 1972, 1, 97-119.

Bohm, Peter. "Environmental Taxation and the Double Dividend: Fact or Fallacy?" Unpublished manuscript. Dept. of Economics, U. of Stockholm, Stockholm, Sweden, Oct. 1995.

Bovenberg, Lans A. and Ruud A. de Mooij. "Environmental Levies and Distortionary Taxation," American Economic Review, Sept. 1994, 94, 1085-89.

Browning, Edgar K. "The Diagrammatic Analysis of Multiple Consumption Externalities," American Economic Review, Sept. 1974, 64, 707-14.

Gilder, Lawrence H., Ian W.H. Parry, and Dallas Burtraw. "Revenue-Raising vs. Other Approaches to Environmental Protection: The Critical Significance of Pre-Existing Tax Distortions," Resources for the Future Discussion Paper 96-24, June 1996.

Gilder, Lawrence H. "Environmental Taxation and the "Double Dividend:" A Reader's Guide," NBER Working Paper Series, Working Paper No. 4896, Oct. 1994

Harberger, Arnold C. "The Principles of Efficiency, The Measurement of Waste" American Economic Review, Jan. 1964, 54, 58-76.

Mirrlees, J.A. "Optimal Tax Theory: A Synthesis," J. Public Economics, Nov. 1976, 6, 327-58.

Ng, Yew-Kwang. "Optimal Corrective Taxes or Subsidies When Revenue Raising Imposes an Excess Burden," American Economic Review, Sept. 1980, 70, 744-51.

Parry, Ian W.H. "Pollution Taxes and Revenue Recycling," J. of Environmental Economics and Management.

Proost, S. And D. Van Regemorter. "The Double Dividend and the Role of Inequality Aversion and Macroeconomic Regimes," International Tax and Public Finance, 1995, 2, 205-17.

Randall, Allan and J. Stoll. "Existence Value in a Total Value Framework," In R. Rowe and L. Chestnut (editors), *Managing Air Quality and Scenic Resources at National Parks and Wilderness Areas*. Westview Press, Boulder, CO, 1983.

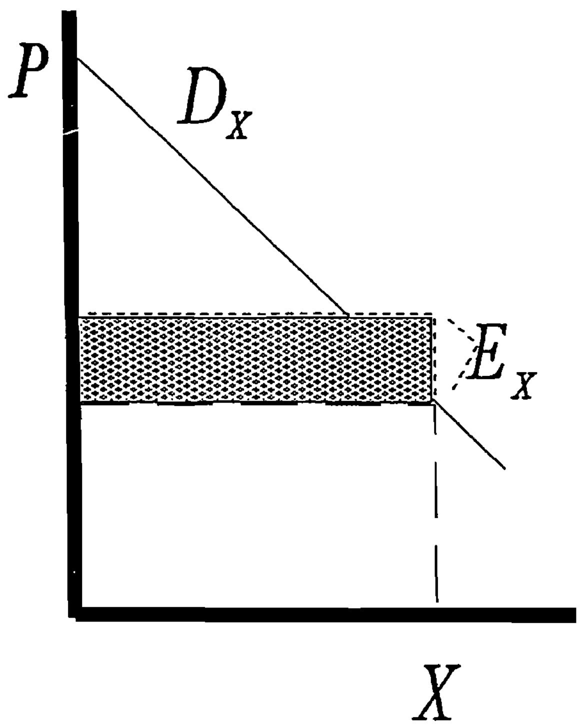
Sadka, Efraim. "On the Optimal Taxation of Consumption Externalities," Quarterly J. of Economics, Feb. 1978, 92, 165-74.

Sandmo, Agnar. "Optimal taxation: An Introduction to the Literature," J. Public Economics, June 1976, 6, 37-54.

Sandmo, Agnar. "A Note on the Structure of Optimal Taxation," American Economic Review, Sept. 1974, 64, 701-6.

Wildasin, David E. "On Public Good Provision with Distortionary Taxation," Economic Inquiry, April 1984, 22, 227-43.

Figure 1



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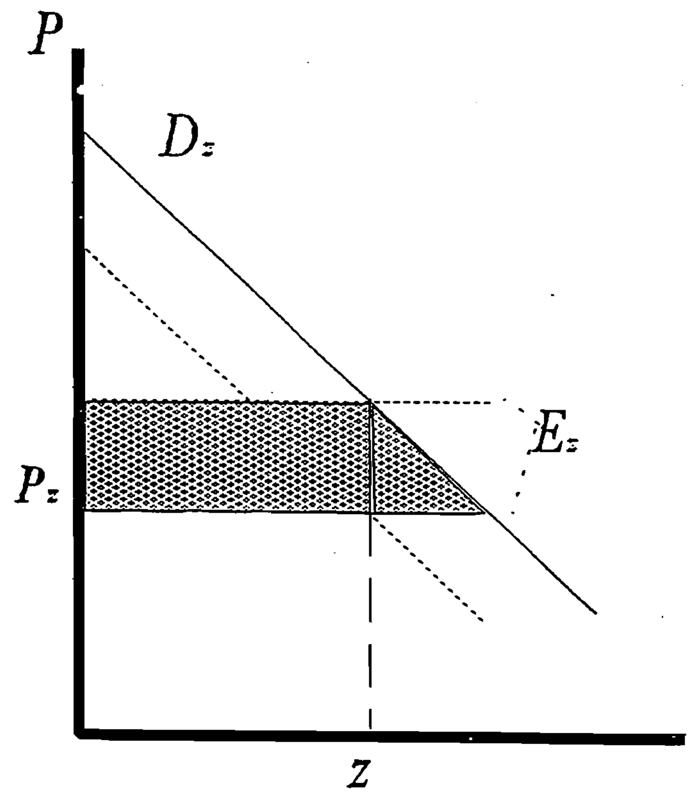

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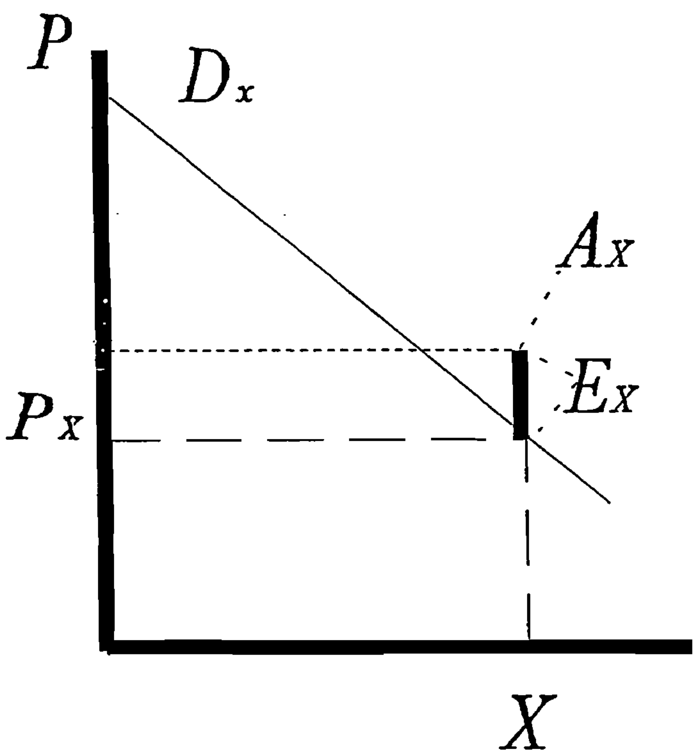

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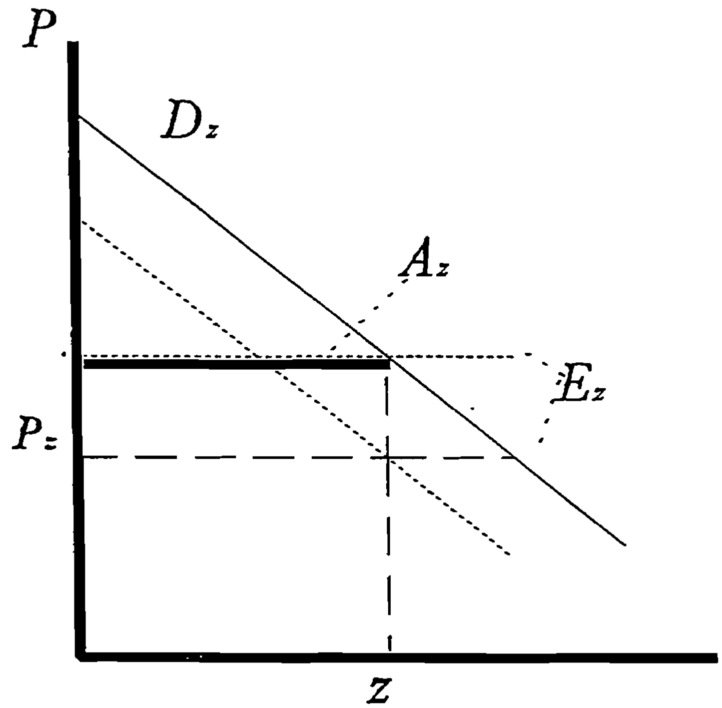

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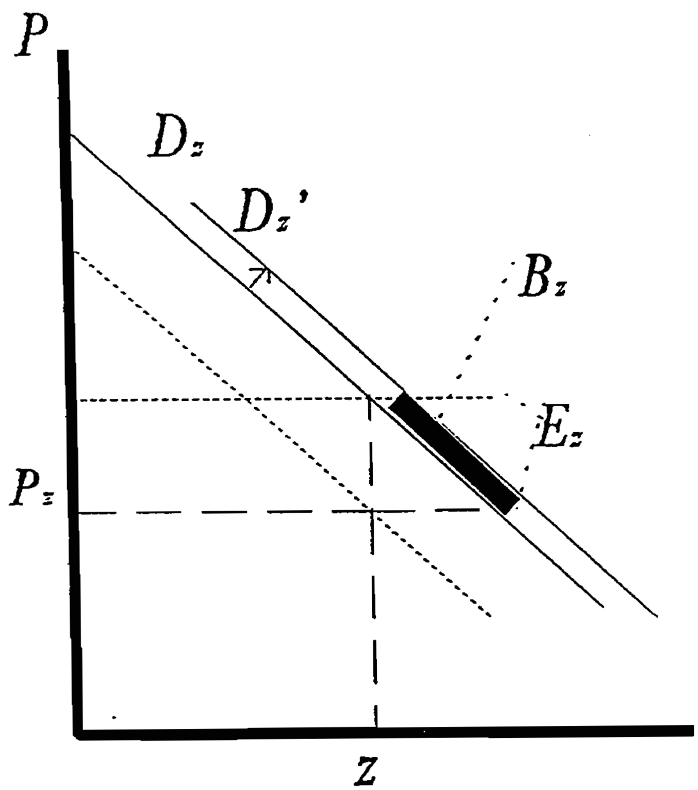

Figure 6