# WESTERN REGIONAL RESEARCH PUBLICATION

W-1133

Benefits and Cost of Resource Policies Affecting Public and Private Land

Papers from the Annual Meeting
Wailea, Island of Maui, Hawaii, February 25-27, 2004
Seventeenth Interim Report
October 2004

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W-1133 Project Objectives:

1. Estimate the Economic Benefits of Ecosystem Management of Forests and Watersheds
2. Estimate the Economic Value of Changing Recreational Access for Motorized and Non-Motorized Recreation
3. Calculate the Benefits and Costs of Agro-Environmental Policies
4. Estimate the Economic Values of Agricultural Land Preservation and Open Space

W-1133 Participating Institutions:
AL, AZ, CA-A, CA-B, CA-D, CO, CTS, GA, IA, KY, LA, ME, MD, MA, MI, NH, NYC, ND, OH, OR, PA, TX, UT, WA, WVA, WY

# Introduction

The meetings held jointly with the Western Regional Science Association (WRSA) were a great time and widely attended. My deepest thanks to Dave Plane for helping our sessions run smoothly. I also must thank my fellow officers, Steve Shultz (president-elect), Klaus Moeltner (acting secretary and vice president-elect) and Ron Fleming (secretary-elect). I must further thank Don Snyder and Fen Hunt for their work as administrative liaisons.

A subset of presented work is provided here as proceedings papers. The presentations were quite interesting, particularly those given at the joint sessions. I believe that the paper-discussant model is quite fruitful for generating valuable feedback for researchers. Note that contact information is included for those interested in making inquiries about ongoing research or papers.

I was honored to be part of the 2004 W-1133 meetings in Maui.

Sincerely,

Donald M. McLeod
University of Wyoming

## W-133 Meeting Attendees – Maui – 2004

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## 2004 W-1133 Meeting Schedule

Wednesday 2/25

Reception with WRSA 4-7pm

Thursday 2/26

*Moderator: Donald McLeod*
The Norton-Leavitt BLM Wilderness Agreement: Economic Benefits to ATV Users and Economic Losses to Wilderness Users
Paul Jakus, Utah State University

An Efficient Experimental Design To Estimate Welfare Measures For A System Of Recreational Areas. Kimberly Rollins and Diana Dumitras University of Nevada, Reno

Public and Hunter Trade-offs between Deer Populations and the External Effects of Deer, Kristy Wallmo, Economist, NOAA Fisheries, Frank Lupi, Associate Professor, Michigan State University, Ben Peyton, Professor, Michigan State University, Peter Bull, Research Associate, Michigan State University

*Moderator: Frank Lupi*
The Value of Lost Fishing Opportunities at Lake Sakakawea from USACOE Management of the Missouri River: Combining Expenditure Distribution & Travel Cost Models. Steve Shultz, North Dakota State University

Modeling Preference Asymmetries in Stated Preference Data: An Application to Rural Land Preservation Robert J. Johnston, University of Connecticut, Kelly L. Giraud, University of New Hampshire

Do Forest Fires Affect Property Values: An Application of the Hedonic Property Model in Colorado. John Loomis and Adam Oren, Colorado State University.

*Moderator: Steve Shultz*
The Effects of Questionnaire Formats on Elicited Preferences and Values in Stated Preference Experiments
John P. Hoehn, Frank Lupi, and Michael D. Kaplowitz Michigan State University

Size Biased Sampling Corrections for Random Utility Models. J. Scott Shonkwiler and Klaus Moeltner University of Nevada

Alternative Strategies for Incorporating Weak Complementarity into Consumer Demand Systems
Roger von Haefen, University of Arizona

BUSINESS MEETING

Introduce Don Snyder, Utah State University, W-1133 Administrative Advisor

CSREES, USDA Budget and Outcomes Criteria Issues: Fen Hunt, W-1133 & CSREES Liaison

Friday 2/27

--------------------**Joint WRSA and W-1133 Session**--------------------------------------------

***Measures of Urban and Rural Property Attributes with Associated Values***

*Chair: Donald McLeod, University of Wyoming*

*"Hedonic Housing Price Analysis with Superfund Sites: The Problem of Bundled Risk Correlates in Urban Industrial Zones."
B. James Deaton and John P. Hoehn, Michigan State University Lansing, MI

DISCUSSANT: Klaus Moeltner, University of Nevada Reno

*"Neighborhood Formation with Fixed Space Constraints: Misspecification of amenity estimates when households are highly heterogeneous."
Michael C. Farmer. Georgia Institute of Technology. Atlanta, GA

DISCUSSANT: Paul Jakus, Utah State University, Logan, UT

"Does ownership matter? Examining the relationship between property values and privately and publicly owned open spaces, streams and wetlands."
Noelwah Netusil, Reed College, Portland, OR.

DISCUSSANT: Donald McLeod, University of Wyoming, Laramie, WY

*"The Opportunity for a Private Market for Farmland Preservation."
Jeffrey H. Dorfman, Bethany Lavigno, John C. Bergstrom and Barry J. Barnett, The University of Georgia, Athens, GA.

DISCUSSANT: John Loomis, Colorado State University, Fort Collins, CO

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*Moderator: Michael Farmer*
"Endangered Species and Data Collection Issues."
Earl Ekstrand, U.S. Bureau of Reclamation.

Consumer Preferences for Locally Made Specialty Food Products Across Northern New England.
Kelly L. Giraud, University of New Hampshire, Craig A. Bond, University of California, Davis, Jennifer Keeling, University of California, Davis

A Watershed Analysis of the Implicit Values of Water Quality and Ecological Diversity
P. Joan Poor, Keri Pessagno and Robert Paul, St. Mary's College of Maryland

*Moderator: Donald McLeod*
Prescribed vs. Perceived Vehicle Costs in Models of Recreation Demand,
Danielle Hagerty, Klaus Moeltner University of Nevada, Reno

Continuous Econometric Models Incorporating Endogenous Stratification and Truncation: An Application to Travel Cost Modeling"
Jeffrey Englin and Laura Nalle University of Nevada, Reno

The Value of Lost Time in Recreation and Work.
Phil Wandschneider and Jon Yoder Washington State University

# Preserving Farmland through Private Markets and Dedicated Funding

Jeffrey H. Dorfman, Bethany Lavigno, John C. Bergstrom, and Barry J. Barnett
The University of Georgia

February 27, 2004

## Abstract

Private land preservation programs are common for rainforest, wetlands, and other ecologically sensitive lands. Farmland preservation, in contrast, has been dominated by government run programs and competition for government funding. Can a privately run farmland preservation program be successful? Surveys of farmers and citizens in Georgia suggest the answer is yes. Results of a random utility model based on a large-scale farmer survey are used to estimate a supply of farmland for preservation by both private and state-run programs. Citizen survey responses are then used to estimate the funding available for such programs, creating an estimate of the demand-side of the farmland preservation market. Results show that voluntary financial support would be sufficient to fund the program at a rate of 8,000-20,000 acres per year depending on the program implemented. Further, the results also indicate that dedicated-source (mandatory) funding for a government-run farmland preservation program would also yield a successful program, with potential protection rates of approximately 40,000 acres per year. All the approaches studied would greatly increase the rate of farmland preservation in Georgia and the rest of the U.S.

**Preserving Farmland through Private Markets and Dedicated Funding**

## Introduction

Farmland preservation is a hot topic in many states as growth pressures bring demand for development of farmland in rural areas. A number of government programs exist for farmers to realize some monetary gains from preserving their farmland in its undeveloped state (either in cash, tax credits, or tax deductions), but only a few scattered private programs are in existence. Transferable development rights (TDR) programs do preserve farmland using private (developer) money, but government is still heavily involved in the program (by awarding development-related benefits at another location in exchange for the farmland preservation. About twenty states fund farmland preservation programs, on top of federal programs funded through USDA.

In contrast, open space and environmentally sensitive lands are being preserved worldwide using private funding. Organizations such as The Nature Conservancy buy the land or the development rights, guaranteeing the permanent protection of the land in an undeveloped state. Such programs have not been particularly directed at farmland, instead focusing on water and habitat related concerns. These programs tend to use exclusively or mostly private money, either from citizen members or corporate donations. The extent of governmental involvement is generally limited to the income tax deduction awarded to the donors in recognition of their charitable donation.

The dichotomy between the major funding mechanisms for preservation of "natural" lands and farmland raises an inherent question about the reasons for the difference. Does farmland preservation funding come mainly from government sources because there is little private, voluntary support for it? If there is no voluntary, private support for farmland preservation, one could rightly question the wisdom of governmentally funded programs directed toward such efforts. Alternatively, it may be that farmland preservation programs make economic and welfare-enhancing sense. Americans may be willing to voluntarily fund farmland preservation, but just haven't been asked.

This paper presents estimates of farmers' willingness to supply farmland for preservation and citizen's willingness to fund farmland preservation. On both sides of the farmland preservation marketplace, we will present empirical evidence for both private and government-run programs. The results will demonstrate that both methods of financing and running a farmland preservation program are viable and could be operated on a much larger scale than is currently being done.

## Methodology

To answer these questions, surveys of Georgia farmers and citizens have been conducted. A mail survey was conducted of Georgia farmers, with surveys sent to 1250 farmers who owned a minimum of 300 acres of land. The farmers were randomly selected by the Georgia Agricultural Statistics Service and all surveys were fully anonymous. After a re-mailing of the survey, we received total responses from 497 farmers (40.74% adjusted for bad addresses) with 390 being usable. The survey had six versions. One-half of the surveys offered farmers a state-run program while the other half offered a privately run program. In addition, three different levels of compensation where offered in exchange for the development rights on 100 acres of land. The question posed to farmers took the form of a yes/no dichotomous choice question. The private version was worded as follows:

> "A private organization in your county is purchasing development rights to farmland in order to permanently protect farmland from development. This group would like to buy the development rights to 100 acres of your farm. You could farm exactly as you do now, and could still sell the land to another farmer, just not to someone who wants to develop it (for houses or businesses). In exchange for the development rights to 100 acres of your farm, the group is offering $3,000 per acre. Would you agree to this transaction?"

The only difference between the above question and the public program version is the replacement of "A private organization in your county" with "The State of Georgia," and "group" by "State." Farmers were offered bids of $1,500, $3,000, or $5,000. A follow-up question then offered a second bid amount based on the answer to the first question, higher for those who answered no and lower for those who answered yes. These follow-up questions looked like the following examples:

> "If yes, would you have sold the development rights for $2,000 per acre?"
> "If no, would you have sold the development rights for $4,000 per acre?"

These second offered amounts were $1,000 and $2,000 for the initial $1,500 bid, $2,000 and $4,000 for the initial $3,000 bid, and $4,000 and $7,500 for the initial $5,000 bid. Along with the willingness to sell questions, farmers were also asked for some standard demographic information and basic details about their farms. The survey was pre-tested with surveys mailed to 252 Georgia farmers that showed little need to adjust questions before the full survey.

The "citizen" surveys were conducted by the University of Georgia Survey Unit who performed a random phone survey of 500 Georgia citizens in fall 2003. These questions were also pre-tested on randomly selected people at Georgia Square Mall in Athens, on the street in downtown Athens, and on the street in downtown Conyers. Participants were asked for basic demographic information, including membership in outdoor and environmental groups and then one of the following three farmland preservation willingness to pay questions:

> "A group of people in your county is forming a private farmland preservation organization. Each member will pay annual dues of $20 for the next five years (and can contribute additional money). All the money will go towards permanently protecting farmland in your county from being developed. The group's goal is to be able to preserve 100 acres of farmland per year. Would you join this group and make the contribution? Yes [·] No [ ].
>
> The State of Georgia is going to sell a new license plate to fund a farmland preservation program. The license plate will have a small picture of a pasture and barn. The tag will cost an annual payment of $20 ($50) in addition to the standard car tag fees for the next five years. The money will all be committed to farmland preservation programs in your county, with a goal of preserving 100 acres of farmland per year. Would you buy one of these license plates? Yes [ ] No [ ]
>
> The State of Georgia is considering holding a referendum this June to begin a dedicated-funding farmland preservation program. If the referendum passes, every taxpayer would pay an additional annual payment of $50 on his/her state income taxes for the next five years (whether they voted for the program or not). All the money would go toward farmland preservation in your county with a goal of preserving 100 acres of farmland per year. Would you vote in favor of this program? Yes [ ] No [ ]"

Note that the voluntary program question was only posed with a program price of $20 per year, the state referendum questions was only posed with a program price of $50, and the optional license plate question was asked with two different prices ($20, $50). This was due to an error by the phone survey unit. Still, by including these three formulations, we can statistically identify the different demands for farmland preservation based on public vs. private programs and voluntary vs. mandatory funding mechanisms. If we only asked public-voluntary and private-mandatory, we could not be sure whether different support levels were due to public vs. private or voluntary vs. mandatory.

## Results

Raw responses from both surveys are presented in tables 1 and 2. The farmer survey responses in table 1 show considerable willingness to sell development rights and preserve their farmland, even at fairly low offer prices. Slightly more than 8000 farmers in Georgia own a minimum of 300 acres (although many of these farms are in South Georgia where growth pressures are less intense). Thus, some quick calculations reveal that a reasonably large amount of farmland is potentially available for preservation

The raw data shown in table 1 was used, along with a set of explanatory variables, to estimate an ordered response probit model of farmland owners' willingness to accept (WTA) for the development rights to their farmland. This model is built from assuming a random utility model as the explanation for farmers' answers to the survey questions on accepting an offer to preserve their farmland. Variables included in the random utility model were: age, gender, education level, income, percent of income derived from farming, years of experience in farming, number of acres owned, number of acres farmed, and regional indicators for North and Central Georgia (leaving South Georgia as the base). Education level was expressed as a number from 1 (some high school) to 6 (graduate degree). Percent of income derived from farming was expressed as a number from 1 (less than 25%) to 4 (more than 75%). For gender, males were denoted by 1, females by 0, so in the model results the coefficient represents the change in willingness to accept amount for men measured relative to women. Also, for interpretation, WTA values (the unobserved dependent variable) are expressed in $1000's, as is income; acres farmed and acres owned are expressed in 100's. Separate models were estimated for the private and state-run programs.

Estimated coefficients, statistical precision measures, and goodness-of-fit measures are shown in table 3. These models performed reasonably well, with an impressive number of variables having statistically (and economically) significant coefficients. In particular, gender made a huge difference and its effect differed across the two models. Men's WTA is estimated to be $2,176 higher when faced with a state program relative to a private program, which suggests an enormous advantage for private programs in terms of the easement acquisition cost. In the private program model, education level is significant and important economically, with WTA dropping by $212 per acre for each additional level of income. Education level is insignificant in the state program model, suggesting that higher education makes one more sympathetic toward a private farmland preservation group, but not toward the state government even when it is trying to achieve the same end. Interestingly, as the percent of income from farming increases, landowners demand more money for their development rights in the private program model, but less in the state program model (although this coefficient has marginal statistical significance). The number of acres owned and farmed makes little difference in either model. Finally, the regional variables are very important with WTA values rising as one moves north in the State of Georgia, which makes sense based on actual observed land values.

Willingness to accept estimates were constructed from the estimated order response probit models. Median willingness to accept values are shown in table 4. These values appear to be completely in line with the actual market value of the development rights for land in Georgia, suggesting that farmers both know the value of their development rights and expect close to full compensation for selling those rights. Given the distribution of willingness to accept values, supply curves for preserved farmland were also constructed, one for the private program and one for the state program. These supply curves are shown in figure 1.

The citizen survey responses are shown in table 2. First examining the two columns on the left, shows people have a very impressive willingness to pay for farmland preservation at a reasonable cost. Faced with a $20 annual price to help preserve farmland, 61% of the survey respondents overall were willing to pay to preserve farmland in their county by joining a private, voluntary organization. For the license plate program, a public, voluntary program there was 56%, support at a $20 price and 36% support at a $50 price. Finally, a very impressive 60% would vote in favor of a mandatory publicly administered program with an annual cost of $50 in additional state taxes. The increase in support for the mandatory program (relative to voluntary-public) suggests that people are more willing to support the concept when they know that others will have to pay as well. This could be because they think a mandatory program will be better funded and hence more successful, or that a mandatory program carries more of a connotation of worthiness. Also, they may support the mandatory program to avoid free-rider problems.

## Potential Markets

The results don't really translate into a region-wide demand curve for farmland preservation, but more properly an ability-to-pay curve representing the number of acres, which can be afforded by the program at different prices. These ability-to-pay curves do slope downward like a demand curve because at higher prices fewer acres can be preserved with the same amount of money. They can also be thought of as budget constraints for such farmland preservation programs. The voluntary program results were translated into such a budget constraint using the number of households in Georgia to compute the total funding available as price multiplied by support rate multiplied by number of households (that is, survey respondents were assumed to be representing their entire household). The license plate results were also computed using the number of households in Georgia (in place of registered vehicles due to skepticism over whether a household would pay additional fees for each car owned). The state referendum results lead to a funding amount equal to the price ($50) multiplied by the number of tax returns filed, since all taxpayers would have to pay the additional tax. These budget constraints are displayed graphically on figures 2 and 3 for the private and public markets, respectively. In the figures, funding was computed on the basis of a five year period, which was the length of the program described to the citizens in the phone survey.

Both figures clearly show the potential for farmland preservation markets as described here. The private, voluntary market could preserve 45,000 acres of farmland in Georgia in a five year period. The license plate program could preserve 37,000 (57,000) acres at $20 ($50) per year for the special plate. The mandatory public program could preserve 180,000 acres given its $50 annual dedicated tax funding, even after accounting for a higher per acre cost need to acquire this many easements. This would be a substantial acreage to protect in five years. Currently, the total farmland acreage preserved in the U.S. is between 1 and 2 million acres (data is sketchy).

## Conclusions

Given the limited funding that has been available from state and federal sources, private funding of farmland preservation could be the force that allows the farmland preservation movement to gain momentum and preserve a quantity of land large enough to make a significant impact on the future of the rural landscape.

Performing the (bold) extrapolation of these results based on a single state, suggests that private, voluntary funding could be sufficient to preserve over 300,000 acres of farmland per year nationally. A mandatory, dedicated funding program at the cost level assumed here, and with development rights costing the same as estimated here, could potentially protect approximately 1.2 million acres of farmland per year (almost as much as has been preserved nationally to date. While this number clearly has huge uncertainty attached to it, it certainly provides an incentive to continue with investigation of both private and public, dedicated-funding farmland preservation programs.

Further research is also needed to determine people's preferences across the type of farmland preservation program and the sensitivity of those preferences across personal characteristics. While the early results suggest that private, voluntary programs can still generate sufficient funds to administer a farmland preservation program of significant size, the funding gap between private and public (mandatory) programs is very large, and cannot be easily ignored.

Another potential topic is to compare the fiscal cost of a governmental program for farmland preservation to its fiscal benefits. Repeated studies have consistently shown that farmland provides a fiscal surplus to local governments, while residential development is a fiscal drain (American Farmland Trust, 1992; Dorfman, et al., 2002). Given the willingness of farmers to sell their development rights for figures as low as $1,000 per acre, it is likely that in some cases, a local government would be better off paying the $1,000 than letting that acre become the site for residential housing.

Finally, perhaps the most important result discovered so far is the potential for dedicated-source public funding. If the societal support shown here is anywhere near reality (when the votes count), the pace of farmland preservation could be accelerated tremendously by taking advantage of the concept's popularity to design a designated-source of funding.

## REFERENCES

Ahearn, M., A. Alberini, J. Bergstrom, K. Boyle, L. Libby, C. Patterson, and L. Racevskis (2000). *Improved Information in Support of a National Strategy for Open Land Policies: A Review of Literature and Report on Research in Progress*. Paper presented at the 24th International Conference of Agricultural Economists. Berlin.

American Farmland Trust (1992). *Does Farmland Protection Pay? The Cost of Community Services in Three Massachusetts Town*s. The Massachusetts Department of Food and Agriculture.

Dorfman, Jeffrey H., Dawn L. Black, David H. Newman, Coleman W.Dangerfield, Jr., and Warren A. Flick (2002). *The Economic Costs of Development for Local Governments*. Report #10 (Center for Forest Business, The University of Georgia). Available on the internet at www.forestry.uga.edu/warnell/pdf/cfb/EcCost.pdf.

Table 1. Farmer Willingness-to-Sell Responses

| Payment Agency | | Initial Amount Offered for the Development Rights | | | |
|---|---|---|---|---|---|
| | | $1500 per acre | $3000 per acre | $5000 per acre | Total |
| Private | | 58 | 56 | 63 | 177 |
| | Yes, Yes | 8 | 7 | 13 | 28 |
| | Yes, No | 12 | 12 | 13 | 37 |
| | No, Yes | 8 | 9 | 7 | 24 |
| | No, No | 30 | 28 | 30 | 88 |
| State | | 63 | 69 | 81 | 213 |
| | Yes, Yes | 7 | 13 | 20 | 40 |
| | Yes, No | 10 | 19 | 22 | 51 |
| | No, Yes | 12 | 10 | 14 | 36 |
| | No, No | 34 | 27 | 25 | 86 |
| Totals | | 121 | 125 | 144 | 390 |

Table 2. Citizen Willingness to Support Farmland Preservation

| | $20 Annual Cost | | $50 Annual Cost | |
|---|---|---|---|---|
| | Percent Supporting | Total Responses | Percent Supporting | Total Responses |
| Private, voluntary program | 60.8% | 142 | -- | -- |
| Public, voluntary program | 56.3% | 71 | 34.2% | 79 |
| Public, mandatory program | -- | -- | 60.5% | 152 |

Table 3. Ordered Response Model Estimates

| | Private Program Model | | | State Program Model | | |
|---|---|---|---|---|---|---|
| **Variable** | **Coeff. Est.** | **Std. Error** | **p-Value** | **Coeff. Est.** | **Std. Error** | **p-Value** |
| Constant | 3.4908 | 0.8085 | <0.0001 | 3.3375 | 0.7730 | <0.0001 |
| Age | 0.0162 | 0.0115 | 0.1625 | -0.0231 | 0.0095 | 0.0162 |
| Gender (M=1) | -0.6292 | 0.3470 | 0.0716 | 1.5466 | 0.4756 | 0.0013 |
| Education | -0.2121 | 0.0811 | 0.0098 | 0.0210 | 0.0979 | 0.8301 |
| Income | 0.0053 | 0.0022 | 0.0145 | -0.0067 | 0.0021 | 0.0014 |
| % Income | 0.3842 | 0.0844 | <0.0001 | -0.1035 | 0.0748 | 0.1680 |
| Yrs. Exp. | 0.0091 | 0.0080 | 0.2558 | 0.0356 | 0.0086 | 0.0001 |
| Acres Owned | 0.0037 | 0.0029 | 0.2045 | 0.0003 | 0.0097 | 0.9753 |
| Acres Farmed | -0.0115 | 0.0078 | 0.1432 | 0.0117 | 0.0101 | 0.2503 |
| North Region | 0.5070 | 0.2466 | 0.0414 | 1.6173 | 0.2310 | <0.0001 |
| Central Region | 0.4294 | 0.2163 | 0.0488 | 0.7439 | 0.1856 | 0.0001 |

Private Program Model: n = 176 log-likelihood = -504.06; correct classifications = 29.0%

State Program Model: n = 213 log-likelihood = -501.89; correct classifications = 36.

Table 4. Willingness to Accept Estimates

| Private Model Results | WTA Estimate | Std. error |
|---|---|---|
| WTA at the median x | $4,527 | $718 |
| Median WTA | $4,988 | |
| North WTA at median x | $5,034 | |
| Central WTA at median x | $4,956 | |

| State Model Results | WTA Estimate | Std. error |
|---|---|---|
| WTA at the median x | $4,287 | $858 |
| Median WTA | $4,780 | |
| North WTA at median x | $5,904 | |
| Central WTA at median x | $5,030 | |

Figure 1. Estimated Supply for Farmland Preservation Programs

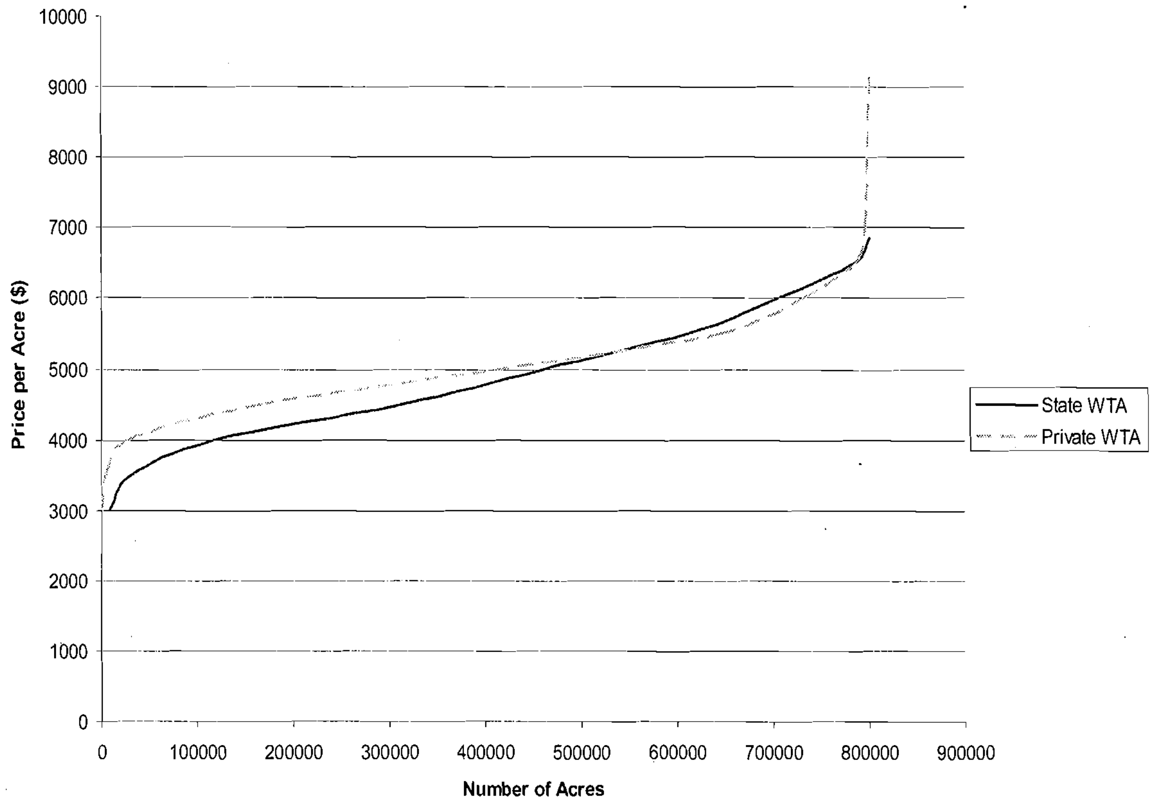


Figure 2. Estimated Public Farmland Preservation Market

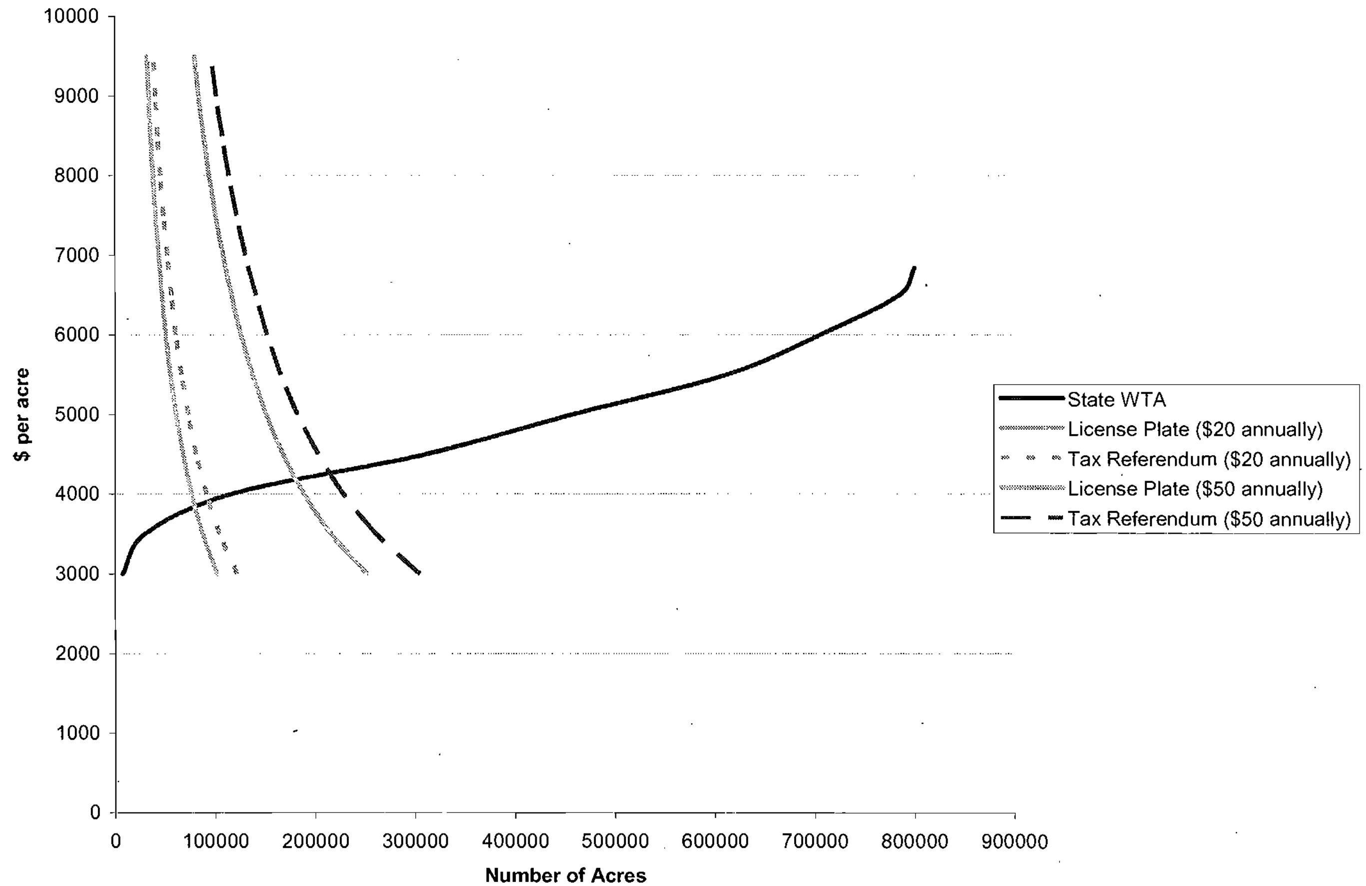
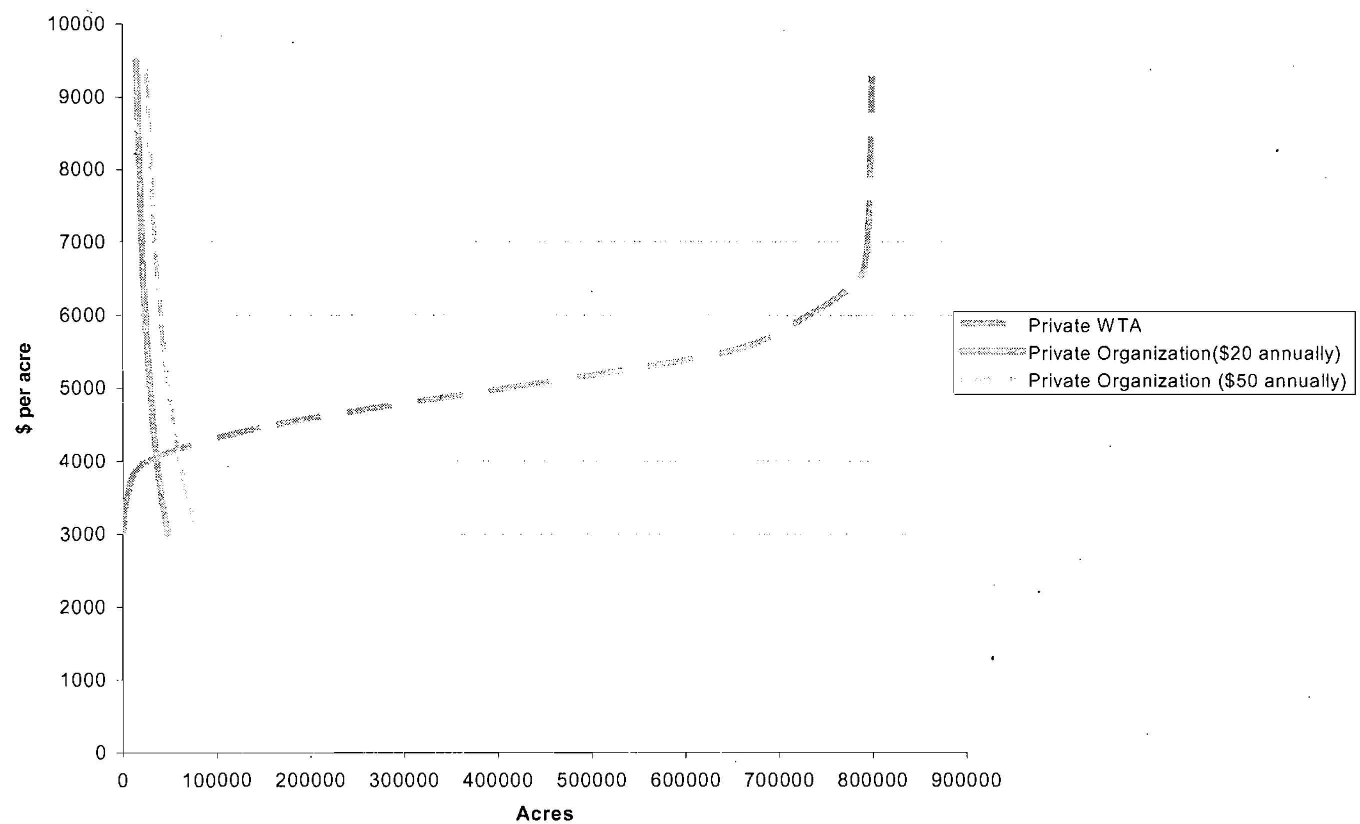


Figure 3. Estimated Private Farmland Preservation Market



# VALUING THREATENED AND ENDANGERED FISH: A COMPARISON OF SURVEY DESIGN

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**Acknowledgments:** This work has benefited from discussions with John Loomis of Colorado State University and Patty Champ of the U.S. Forest Service.

**Disclaimer:** The opinions expressed here do not necessarily reflect the policy or views of the U.S. Bureau of Reclamation.

This work was funded in part by the U.S. Bureau of Reclamation's Science and Technology Research Program, Number A10-1541-3050-158-01-01-8 and A10-1742-8052-350-01-0-1-1.

## Summary

Valuing threatened and endangered species (T&E) requires economists to estimate values using a hypothetical market method within a survey. A survey is presented to a sample of the public to determine a willingness to pay (WTP) for protection of the species of interest. It is necessary that the sample represent the population relative to age, education and other demographics. However, acquiring a representative sample, along with administering the survey, is costly in time and money.

This research considers an alternative method of selecting a sample, utilizing a consumer panel maintained by an international marketing firm. The marketing firm maintains a list of potential respondents that can be accessed regionally or nationally. Consumer panels are balanced relative to the demographics of the population in question. Using the consumer panel is less expensive and respondents are more used to the idea of responding to surveys, since they occasionally receive surveys.

A consumer panel was asked to respond to a survey valuing T&E fish species in the four corners region of the U.S. In this region nine fish species are listed as either threatened or endangered with critical habitats along six rivers designated by U.S. Fish and Wildlife Service. These habitats are affected by diversions and other operations on the rivers, with designation of critical habitat altering management of diversion facilities.

The T&E survey was originally administered to the public in 1997 using the traditional approach of acquiring a list of potential respondents, sending the survey and following up with reminder postcards and a second mailing.[1] The original approach yielded 718 responses, a 54 percent response rate. Analysis yielded a WTP of about $279 per household.

The marketing firm readministered the survey in 2001 using the consumer panel. The survey used the same text and questions as in the original survey and other pertinent information along with the same map, providing a description of the species and critical habitat. The same methodology was used to analyze the resulting data.

This resampling resulted in 432 usable responses (a 61 percent response rate) from respondents across the U.S. Analysis of these responses used the same variables analyzed in the original study. Preliminary analysis of this retest resulted in a WTP of $148.

The original analysis estimated a WTP of $279, significantly different from the second estimate of $148. Confidence intervals were constructed for these two estimates. For the original study, with a mean of $279, 95 percent confidence intervals ranged from $191 to $548. The 95 percent confidence intervals for the second analysis ranged from $114 to $212, around the $148 mean.

[1] Ekstrand, E. R. And J. Loomis. Incorporating Respondent Uncertainty When Estimating Willingness to Pay for Protecting Critical Habitat for Threatened and Endangered Fish. Water Resources Research, 34:11, pages 3149-3155. November 1998.

While the means differ by $131, the confidence intervals overlap. Analysis shows that the means for this model and the other two models, are significantly different.

The results of the study have policy implications for management of rivers that provide critical habitat. Both results show that the public values these habitats and preservation of these species with the WTP values that are significantly different from zero and are substantial, considering the numbers of households represented by these samples.

## Introduction

Decisions involving use of public lands often are based on economic benefit-cost analysis. Public lands include many resources, such as threatened and endangered species (T&E), free-flowing rivers, and wilderness areas that are valuable to society but do not have market prices. Without market prices, benefit estimations for these resources are difficult to acquire. But because society often recognizes real opportunity costs resulting from protecting these public resources, it is necessary to recognize the economic benefits and not just the economic costs.

Though difficult to estimate, these benefits are often large, relative to costs. For example, numerous studies of T&E species have shown large economic benefits due to preserving these species (Loomis and White 1996). Values per household for some of these resources may be low, but large aggregate values result due to millions of households throughout the U.S. Since all households can simultaneously enjoy the benefits of knowing these species still exist. But due to the dispersed nature of public goods, little incentive exists for beneficiaries to be actively engaged in the policy process.

Measuring benefits using willingness to pay (WTP) is the currently accepted norm among Federal agencies for benefit-cost analysis (U.S. Water Resources Council 1983) and natural resource damage assessment (U.S. Department of Interior 1986). Because of the lack of market price, economists have developed a hypothetical market method called the contingent valuation method (CVM) that uses a survey to measure household WTP for public goods, including T&E species. A CVM survey is a standardized and widely used method for obtaining WTP. It involves developing a hypothetical market or referendum as a vehicle by which an individual reveals his or her WTP. The CVM is recommended for use by Federal agencies for performing benefit-cost analysis (U.S. Water Resources Council 1983), for valuing natural resource damages (U.S. Department of Interior 1986), and was upheld by the Federal courts (U.S. District Court of Appeals 1989). A panel including two Nobel laureate economists, an environmental economist and a survey research specialist reviewed the CVM process and, while they felt that this method had limitations, they concluded that CVM can produce estimates reliable enough to be the starting point for administrative and judicial determinations (Arrow et al. 1993).

However, performing a CVM survey is a costly process in both time and money. In order to perform a national survey, a sample of the population of the U.S. needs to be acquired, a questionnaire prepared and printed, mailings executed and the responses tabulated. This has been a limiting factor for agencies with limited budgets. Responses from a consumer panel cost approximately a third as much per respondent, compared to a national survey. Therefore, this research project looked at the possibility of using an alternative approach to a national survey by using a consumer panel that represents a sample of the population of the country.

This comparison is accomplished by using the consumer panel to estimate willingness to pay and comparing it to the results from a previous survey in 1997, both studies valuing the same good (Ekstrand and Loomis 1998). This comparison is to determine whether the results of the consumer panel are similar to the national survey and, if there are differences, identify what the differences are.

## Valuing Nine Fish Species

Ekstrand and Loomis (1998) estimated the willingness to pay for protecting critical habitat for nine fish species in the Four Corners region of the U.S. Using a 1997 national survey, the original study compared the standard single dichotomous choice CVM to alternative modifications that explicitly incorporate respondent uncertainty for the purpose of estimating economic benefits of protecting critical habitat for nine T&E fish species living in the Colorado, Green and Rio Grande river basins. The standard dichotomous choice CVM estimated a value of $268 per household, which was compared to values ranging from $50 to $330, depending on how respondent uncertainty was explicitly incorporated into the dichotomous choice model.

Most individuals are not familiar with many T&E species and have no prior experience paying for species protection. Many individuals realize personal satisfaction from knowing these species exist but have not devoted much time contemplating how much they would pay to protect critical habitat. If they spent the time to reflect on the tradeoffs between household costs and preservation of species, they could refine their preferences. However, the one-shot nature of CVM survey responses may not provide sufficient repetition for generating stable preferences.

While CVM may not provide the opportunity to stabilize preferences through repeat purchasing behavior, respondents may be able to express the level of confidence in their dollar bids, and this information can be incorporated into the statistical analysis. Those individuals who have extensive prior knowledge of the environment or species in question may have well-defined preferences and great certainty in their responses, while those with little or no knowledge may have less-defined preferences and therefore more uncertainty about their answers. Incorporating the stated uncertainty of respondents into the statistical model could improve the estimation and accuracy of the analysis (Manski 1995). For a more detailed discussion on incorporating respondent uncertainty into models estimating willingness to pay for these critical habitat units, please refer to Ekstrand and Loomis (1998).

The intention of this research is to compare the results of the 1997 study with new results estimated using a consumer panel provided by the marketing firm NFO Worldwide (NFO). The original 1997 survey was repeated in the fall of 2001 using samples from NFO consumer panels. NFO Worldwide is a research firm and is a provider of custom, research-based marketing information and Internet-based research. NFO states their consumer panels provide a known base of research about whom detailed information is known. They track 200 different data points on individual's' attitudes, beliefs and behavior each year in about 575,000 U.S. households, in addition to online individuals available via the Web. After receiving the text to be used in the survey from the analyst, NFO prepares the document, mails the survey and tabulates the responses, providing the analyst the results in an electronic form.

Previous studies have shown that there is reliability in the test-restest method for contingent valuation. Loomis (1990) retested after a nine-month wait and found reliable results. Others found similar results as discussed in Whitehead and Hoban (1999). However, we are not aware of any test-retest that used different sources of sampling, as in this study.

## Hypothesis

The models in this analysis followed two of the models in the original analysis. For comparison purposes, we analyzed the data, adjusting the affirmative responses to the WTP question based on the responses to the certainty question. We estimated the standard dichotomous choice WTP using the logit model:

(1) $\text{Prob (YES)} = 1-\{1+\exp [B_0 - B_1 (\$X)]\}^{-1}$

where \$X is the dollar amount the individual is asked to pay and $B_0$ and $B_1$ are the intercept and slope coefficients, respectively.

For each sample, the standard dichotomous choice response was used as the first model to be analyzed, referred to as the SD model. Then two more models were analyzed for each sample, using information contained in the 1 to 10 post-decisional ranking from the respondents regarding their certainty to the WTP response. As in the 1997 study, a YES10 model was created, recording all YES responses as NO responses if the respondent indicated a certainty level less than 10, based on the scale of 1 being very uncertain and 10 being very certain, in addition to all NO responses being coded NO. A YES response with a certainty of 10 was coded as a YES response. Finally, a third model was created for each sample, the YES9 model, which coded all YES responses with certainty of 9 or 10 as a YES and all YES responses with certainty less than 9 and all NO responses as NO.

Our hypothesis in this study is that the comparative models from each sample will have statistically similar WTP values. That is, the WTP from the SD models of each sample will be equal, the WTP from the YES10 model in each sample will be equal, and the WTP from the YES9 model in each sample will be equal.

## CASE STUDY: REVALUING CRITICAL HABITAT FOR NINE T&E FISH SPECIES IN RIVERS OF THE FOUR CORNER STATES

Ekstrand and Loomis, in the original study, estimated values for preserving critical habitat units (CHUs) for nine T&E fish species in the Four Corners region of the U.S. There are six rivers in these water basins. The Colorado River provides habitat for the bonytail chub, Colorado pikeminnow, humpback chub, and razorback sucker. The San Juan and Green rivers contain the Colorado pikeminnow and razorback sucker. The Virgin River in Utah provides habitat for the Virgin River chub and the woundfin. The Loach minnow and spikedace are species in the Gila River in Arizona and New Mexico, while the silvery minnow lives in the Rio Grande River in New Mexico. Designation of critical habitat affects instream flow requirements and alters management of hydropower facilities. See Ekstrand and Loomis for details of these CHUs.

The 1997 study used a random sample of residents throughout the U.S. The 1997 study consisted of a 12-page survey document with a cover letter, plus a map of the region showing the rivers and locations of the critical habitat units along those rivers. The 1997 survey was based on input from three focus groups, one each in Fort Collins, Colorado; Albuquerque, New Mexico; and Phoenix, Arizona. These focus groups led to revisions based on the suggestions and comments of the participants. Following the focus groups, the research team developed a complete mail booklet and survey script used to pretest a small sample of households throughout the U.S. After comments from respondents in the pretest, the final survey document was prepared.

The first section of the survey allowed the respondents an opportunity to reflect on why they might care about the endangered species. The first set of questions asked about the relative importance of federal lands for providing habitat for endangered species versus using resources for extraction and jobs. A five-point Likert scale allowed individuals to agree or disagree with a set of attitude questions to measure how utilitarian the respondents were versus how preservation oriented they were. These responses also provided insight into the responses to the WTP question and have been used in other research (Barrens et al. 1996).

Our CVM survey followed the standard three-element design: (a) portrayal of the resource to be valued, (b) description of the particular mechanism to be used to pay for the resource and (c) the question format used to elicit the respondent's dollar amount of WTP. The resource being valued was the 2,456 miles of CHUs as described in the text and shown on the map. Protection involved habitat improvements, such as fish passageways and bypass releases of water from dams to imitate natural water flows needed by fish. A list of fish species by river was printed in the survey.

Respondents were told that some State and Federal officials thought the costs of the habitat improvements and the restrictions on hydropower were too costly, and that proposals for eliminating CHUs had been put forward. Then the description of the particular mechanism to be used to pay for the resource was provided. Respondents were told the current program could be paid for by the establishment of a Four Corners Region Threatened and Endangered Fish Trust Fund. Efforts to raise funds would involve all U.S. taxpayers contributing to this fund. If a majority of households voted in favor, the fund would maintain CHUs for the nine T&E fish species, to avoid extinction. This would be accomplished through water releases from Federal dams timed to benefit fish and the purchase of water rights to maintain instream flows. The survey stated that, within the next 15 years, three fish species would increase in population to the point they would no longer be listed as threatened species.

However, if a majority of households in the U.S. voted to not approve, then the CHUs shown on the enclosed map would be eliminated. That would mean water-diversion activities and maximum power production would occur, reducing the amount of habitat for these nine fish species, and that, as a result, biologists estimated that four of the nine fish species would most likely become extinct in 15 years.

This information was followed by the question format used to elicit the respondent's dollar amount of WTP, which asked each household how they would vote, considering the price indicated. This referendum format is recommended by the panel on CVM (Arrow et al. 1993). The exact wording on the questionnaire was:

Suppose a proposal to establish a Four Corners Region Threatened and Endangered Fish Trust Fund was on the ballot in the next nationwide election. How would you vote on this proposal?

Remember, by law, the funds could only be used to improve habitat for fish.

1. If the Four Corners Region Threatened and Endangered Fish Trust Fund was the only issue on the next ballot and it would cost your household $_____ every year, would you vote in favor of it? (Please circle one.)

YES NO

The dollar amount, which is blank in this example, was filled in with one of 14 amounts ranging from $1 to $350, randomly assigned to survey respondents. The range was picked such that at the low end, anyone who valued preserving the fisheries protection would very likely indicate they would pay $1-3, while almost no one was expected to pay $350 per year.

On the next page of the survey, respondents were asked to determine how certain they were when answering the WTP question. The wording in the survey was as follows:

2. On a scale of 1 to 10, how certain are you of your answer to the previous question? Please circle the number that best represents your answer if 1=not certain and 10= very certain.

| 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 |
|---|---|---|---|---|---|---|---|---|---|

not certain < - - - - - - - - - - - - - - - - - - - - - - - > very certain

After this section, allowing the respondent the opportunity to express views on these nine fish species, an additional section of the survey document then asked respondents to consider protection of a larger number of species in the Four Corners region of the U.S. In addition to the nine fish species, 53 additional species were identified as T&E species, including fish, birds, mammals and plants. In addition to the critical habitat for the nine fish species, additional land was identified as being critical to the additional species, along with pertinent information related to protection. This section was similar to the section that discussed the program for the nine fish species, ending with a similar request to vote, using a single dichotomous choice question, as used before. This section was accompanied by a separate map indicating the critical habitat units for the 62 species discussed in the section.

The 2001 survey was not as extensive. The section asking the respondent to value the 62 total species was not included, as the scope test was not a concern in this retest. Also, in the 1997 survey, the final section asked the respondent to reply to demographic questions, including residence location, age, occupation, education and income. Because the demographic information on the consumer panel had already been collected by NFO, the demographic section was omitted from the retest survey document. Included in the retest document were the sections giving the respondents an opportunity to reflect on why they might care about endangered species, along with the section describing the resource to be valued, the particular mechanism to be used to pay, and the question used to elicit the dollar amount of WTP. The exact text from these sections of the original survey was used in the retest document, along with the same map that was used in 1997 showing the rivers and critical habitat units for the nine fish species. The exact wording from the original document was used to keep the comparison as close as possible. Using just the subsection of the original survey document allowed the retest document to be four pages, plus the map.

## Survey Results

The 1997 survey was mailed to a random sample of 1,600 households in the U.S., half to the households in the Four Corners states, with proportions based on the relative populations. The other half of the surveys was mailed to households in the other 46 states, again weighted by population. The sample was provided by Survey Sampling, Inc., a company that specializes in providing representative samples and one that has been frequently used by researchers in the past.

The overall design and mailing procedure followed Dillman's (1978) Total Design Method. In addition, a dollar bill was included with the first mailing as a token of appreciation and to increase the response rate. Each individual was sent a personalized cover letter on university letterhead with an original signature. Both the outgoing and return envelopes had a first-class postage stamp affixed to further distinguish the mailing from bulk mail, followed by a reminder postcard. A second mailing was performed (without the $1 bill) to nonrespondents. After adjusting for nondeliverables and deceased, the 718 returned surveys represented a 54 percent response rate, as shown in Table 1.

The 1997 survey overweighted the households in the Four Corners states. This overweighting was adjusted when the data from the subgroups were pooled. The 46-state sample was overweighted in the likelihood function during the analysis.

The 2001 retest used the procedures of the NFO marketing firm, consisting of one mailing with no followup postcard or second mailing to nonrespondents. Even with no followup material, of the 713 surveys mailed, 432 surveys were returned, a 61 percent response rate, as shown in Table 1. The NFO consumer panel members also represented a random sample of the households in the U.S. In this retest, the Four Corners states were not overweighted in the sample, as this retest was interested only in comparing national results to the 1997 study.

Comparison of income levels between the two samples showed similar characteristics. In the original sample, the respondent's average income was about $54,000, while the average for the respondent on the retest study was about $59,000. The consumer price index difference between the two years is about 10 percent, leading to an adjustment of about $5,400 to the 1997 amount, equaling $59,400 in 2001 dollars. When adjusted for inflation, the two incomes are very similar.

Both studies showed diminishing affirmative responses as the dollar-bid amounts increased, as shown in the Table 2. For those receiving the referendum question asking if they would pay $1, over 70 percent responded YES, in both studies. Similarly, the percentage of YES responses decreased as the bid amount increased, as expected by economic theory.

**Statistical Estimation of the Logit Model** In both studies, the printed dollar amount varied across the sample of respondents, ranging from \$1 to \$350. With this variation, the voter-referendum format requires the analyst to statistically trace out a demand-like relationship between probability of a YES response and the dollar amount, using a qualitative response model such as logit or probit (Hanemann 1984). The basic logistic regression model was given in equation (1).

From equation (1), Hanemann (1989) provides a formula to calculate the mean or expected value of WTP, assuming WTP is greater than or equal to zero. The formula is:

(2) $$\text{Mean WTP} = (1/B_1) * \ln(1+e^{B_o}) \text{ where WTP} \geq 0$$

For this analysis, $B_1$ is the coefficient estimate on the bid amount, and $B_0$ is the sum of the estimated constant plus the product of the other independent variables times their respective means for each respondent. The individual means are then averaged to determine an average mean WTP for the sample.

Besides the bid amount, independent variables included income, and proxies for tastes and preferences called PROTECT and PROTJOB. PROTECT was the sum of the answers on the Likert scale from the questions asking about the desirability of protecting plants and animals.

PROTJOB was the sum of the responses to the Likert scale questions related to the rights of business to extract resources and be protected from loss of jobs. As the variable PROTECT was the sum of 4 questions, and the variable PROTJOB was the sum of two questions, PROTECT was divided by 2 so that the coefficient would compare to the PROTJOB coefficient. Also, in the 1997 survey, because the Likert scale asked the respondent to answer 1 for strongly agree and 5 for strongly disagree, the coefficients to these variables would be intuitively reversed. Therefore, each was multiplied by a negative one to reverse the signs of the coefficients. However, in the retest survey, the Likert scale questions asked the respondent to answer 5 for strongly agree and 1 for strongly disagree.

## Statistical Results

Table 3 provides the coefficients for the logit equations for these preliminary results. In the original study, the coefficients on all variables were statistically significant, but in the retest study, the income variable was not significant in the three estimations of WTP. In all estimations, the coefficients for the bid amount (PRICE) were negative and significant, indicating that as bid amount increases, the respondent is less likely to pay. The coefficients for income have positive signs, as incomes increase, willingness to pay increases.

PROTECT generated positive significant signs on these coefficients, showing strong preferences about protecting endangered species, and were more willing to bid higher dollar values. Those respondents with high scores on the PROTJOB variable emphasize employment above T&E protection, and this results in significant negative signs on the coefficients, implying less likelihood of paying to protect the nine T&E fish. In all estimations, the coefficients were significant at the 99 percent level for both variables.

Mean WTP values were calculated using Equation 2, with the resulting values shown in Table 3. The mean WTP for the standard referendum question was \$278 per household (after reweighting due to the oversampling in the Four Corners states) using the original study data,[2] but in the retest estimation, the WTP is \$148. Similar estimations are provided for the other four models in this study with similar results. The values shown in this table for the original study data provide larger WTP values than in the retest data. For example, in the YES10 model, original study data provide a WTP value of \$35, while the retest YES10 WTP is \$21.

Table 3 also shows the confidence intervals for these WTP estimations. In the original published study (Ekstrand and Loomis 1998), 95 percent confidence intervals were estimated using the standard errors of the WTP estimations. In this retest, comparisons of the WTP were performed using the Krinski-Robb (Krinski and Robb 1986) technique and using convolutions method as suggested by Poe et al. (1994). These Krinski-Robb confidence intervals are shown in Table 3. The reader can see that, while there is some overlapping of confidence intervals between comparative models, these overlapping regions are limited in size, leading to the conclusion that the WTP estimations are not statistically similar. This conclusion is verified by the convolutions test. In all tests, the probabilities of the estimations being statistically similar are very small, as shown by the P(0) values in this table.

[2]The original study published a WTP of \$268. In this reestimation of the original data, additional observations were omitted due to missing data. A similar difference occurred for the YES10 model, the published WTP is \$50.

## Conclusions

For the purpose of policy decisions, this retest showed there is significant value in protecting these nine T&E fish species' critical habitat units, as estimates of mean WTP are statistically different from zero, as was concluded in the original study. In this retest, the lowest estimated willingness to pay was $21, with a lower bound of $12. With approximately 100 million households in the U.S., the economic benefits of protecting these habitats is substantial, even with these estimations based on the assumption that all nonrespondents have a WTP of zero.

However, comparison of the two samples does not provide the same WTP estimations in the preliminary analysis of these data, contrary to previous studies performing retests for the same resource. This difference in WTP may be due to the use of the consumer panel or due to other factors.

One reason that the difference in WTP may occur is that the consumer panel does not represent the same population as the original sample. This would imply that either the original sample or the consumer panel was not representative of the population as a whole. However, there appeared to be no evidence that these two samples represented different populations.

Differences in WTP could have occurred because the consumer panel had more experience with expressing WTP for products, especially new products with which they have no experience. The consumer panels receive surveys from NFO on a periodic basis, some of which ask the respondents to value a new product that a manufacturer is interested in selling. Usually, the new products are marketplace products, different from habitat units for T&E species, but the respondents still have some experience in valuing unknown goods.

As this presentation is preliminary analysis of the data, additional testing may reveal more information. Comparisons of the results of the certainty responses may indicate whether the consumer-panel respondents are able to express WTP with more certainty than the sample from the general population.

As shown in Whitehead and Hoban (1999), tests for changes in attitudes can be performed to determine why the WTP for these resources were lower in the 2001 sample. A decomposition of WTP can be conducted to determine whether the difference in WTP is due to differences in factors affecting WTP or the underlying structure of the coefficients of WTP.

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Table 1: Response Rates.

| Sample | Number Mailed | Number Returned | Percent Returned |
|---|---|---|---|
| Original Sample | 1,332 | 718 | 54% |
| Retest Sample | 713 | 432 | 61% |

Table 2: Affirmative Responses for Each Bid Amount.

| | Percent Affirmative | |
|---|---|---|
| $ Bid Amounts | Original Study | Retest Study |
| 1 | 71 | 72 |
| 3 | 84 | 70 |
| 5 | 74 | 68 |
| 10 | 56 | 85 |
| 15 | 58 | 74 |
| 20 | 61 | 59 |
| 30 | 60 | 31 |
| 40 | 57 | 50 |
| 50 | 47 | 54 |
| 75 | 43 | 49 |
| 100 | 50 | 41 |
| 150 | 39 | 37 |
| 200 | 42 | 39 |
| 350 | 39 | 25 |

Table 3: Coefficients, WTP and Confidence Intervals.

| | SD | | YES10 | | YES9 | |
|---|---|---|---|---|---|---|
| Variable | Original | Retest | Original | Retest | Original | Retest |
| Constant | 0.92 | -2.60* | -2.46* | -5.47* | -0.88 | -4.98* |
| Price | -0.0040* | -0.0077* | -0.0070* | -0.0095* | -0.0055* | -0.0094* |
| Income | 0.0099* | 0.0021 | 0.0067* | 0.0056 | 0.012* | 0.0052 |
| Protect | 0.45* | 0.31* | 0.34* | 0.34* | 0.44* | 0.35* |
| Protjob | -0.20* | -0.31* | -0.32* | -0.32* | -0.21* | -0.33* |
| WTP | $279 | $148 | $35 | $21 | $64 | $33 |
| Confidence Intervals** | $191 - 548 | $114 - 212 | $23 - 59 | $12 - 37 | $41 - 114 | $22 - 52 |
| P(0)*** | | 0.010 | | 0.077 | | 0.022 |

*significant at the 99% level
**based on the Krinski-Robb technique, assuming a 95% confidence interval range
***Probability of WTP overlaps between two samples

# Consumer Preferences for Locally Made Specialty Food Products Across Northern New England

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**Consumer Preferences for Locally Made Specialty Food Products Across Northern New England**


## Abstract

Does willingness to pay for local specialty food products in Maine, New Hampshire and Vermont differ? Two food categories are investigated: low-end ($5), and high-end ($20) products. Premia estimates are compared across states and across base prices within states using the Method of Convolutions. Results suggest that the three states of Northern New England have many similarities, but residents in Maine are willing to pay more than residents of New Hampshire for a low-end product. Vermonters and New Hampshirans are willing to pay a higher premia for a $20 over a $5 food item , while Mainers are not.



**Keywords:** local specialty foods, willingness to pay, contingent valuation

## Introduction

The states of Northern New England, Maine, New Hampshire and Vermont, often conjure images of lobsters, blueberries and maple syrup for residents and visitors alike. Indeed, the distinct style of locally grown and produced specialty food items contributes to the economic vitality of the region through trade and tourism. The governments of both Maine and Vermont and the citizens of New Hampshire have recognized these contributions and worked to improve the regional economy, increase local employment and promote agriculture in the area through the implementation of marketing programs for locally labeled produce and specialty foods.

The demand for specialty foods has been especially strong in recent years, and it is estimated that one in five U.S. household can be classified as a medium to heavy consumer of specialty food items (Kezis et al. 1997). However, very little research has been conducted to investigate state-made product preferences for items other than fresh produce, nor has extensive research been done to identify preferences for local goods in the New England region. As such, this paper extends the literature by investigating the preferences of Northern New Englanders for locally produced specialty food products. Following Peat, Marwick, Stevenson, and Kellogg (1990), we define a specialty food to be a value-added, premium priced item that is distinguished in terms of one of more characteristics such as the quality of ingredients, sensory appeal, origin, presentation including branding or packaging, or the product formulation.

The objective of this paper is to address the question of whether Northern New England residents admit preferences that favor state-made specialty goods over imported substitute goods, and if so, what price premium can be supported? In the absence of well defined local product differentiation in actual market data, the question of consumer willingness to pay is answered using the contingent valuation method, treating the state of origin as the distinguishing quality attribute of a homogenous good, and estimating the value of that attribute. The heterogeneity of consumer perceptions across states is discussed, and local price premia, where they exist, are estimated for both a relatively low and high priced specialty food. The premia are then tested for equivalence across states and across goods.

The paper proceeds as follows. The next section discusses local labeling programs and previous literature regarding preferences for locally grown goods. The contingent valuation model is then described, followed by a brief discussion of the survey design. Basic survey results are then presented, including demographics and consumer perceptions of locally grown food products. Next, findings on consumer willingness to pay for local attributes are presented followed by a final section that concludes and summarizes the results.

**Review of Local Labeling Programs Research**

Following the success of state funded local labeling programs in states such as New Jersey ("Jersey Fresh") and Tennessee ("Tennessee Proud"), Govindasamy, et al. (1999), report that as many as 23 states have enacted their own local labeling and marketing campaigns in an effort to increase sales of locally grown or processed food. Several studies have found the consumer loyalty for local products is enhanced by awareness of local goods and state labeling and promotion programs (Wolfe and McKissick 2001, Govindasamy et al. 1998, Jones et al. 1990, Brooker and Eastwood 1989). This indicates that state labeling programs have the potential to successfully differentiate local goods and increase niche market sales if target consumers are exposed to promotional material. In addition, Brooker and Eastwood (1989) find that consumers would prefer not to have a single label used for both fresh and processed local foods; however, only Florida and Tennessee currently use different labels to identify locally processed and locally grown foods (Thomas et al. 2001).

Promotion of state labeled produce and processed goods were found to take several forms, including basic labeling of goods, in store display/signs, and television advertising (Wolfe and McKissick 2001, Govindasamy et al. 1998, Thomas et al. 2001). In two particular studies, in-store taste tests and sampling were demonstrated to be particularly effective methods of promoting local produce and processed foods (Wolfe and McKissick 2001, Kezis et al. 1997). In point of fact, Dietrich (1992) and Kuryllowicz (1990) have found that 70% of all customers in specialty food stores will accept a sample and nearly one-forth will buy the product after sampling. Regardless of the method of promotion, nearly all studies emphasized the need to inform consumers of the superior quality, freshness, uniqueness, and greater diversity of products crafted or grown in the state. Through these advertising campaigns, consumers may be educated to differentiate between local goods and imports , and thus shift their preferences towards locally produced goods.

Successful differentiation causes local brands to be more appealing, and often results in a price premium which can be measured as the consumer's willingness to pay or purchase a good or willingness to switch stores to purchase a local good (Wolfe and McKissick 2001). The majority of studies found that consumers were willing to pay a qualitative premium for fresh local produce, but the percentage of consumers varies by state and commodity studied, and few consumers were willing to switch supermarkets to be able to purchase locally produced fresh produce. The current literature concerning consumer willingness to pay for state labeled produce has focused nearly exclusively on fresh local produce, and on the qualitative question of whether or not a price premium exists. One notable exception is Loureiro and Hine (2002), who quantify differences in willingness to pay across niche market characteristics for fresh potatoes. We extend the current literature by measuring consumer's willingness to pay for locally produced processed or specialty goods, rather than fresh produce, across subregions of Northern New England and compare these premia across states. To do so, we employ dichotomous choice contingent valuation analysis and the method of convolutions.

**The Contingent Valuation (CV) Model**

Following Haneman (1984), we now present the basic binary choice utility model used in this analysis to estimate consumer willingness to pay for local specialty goods. Suppose an individual $\boldsymbol{n}$ is faced with a choice between $\boldsymbol{i}$ (buying the local specialty food product) and $\boldsymbol{j}$ (the non-local specialty food product). Product $\boldsymbol{j}$ costs \$$\boldsymbol{A}$ and product $\boldsymbol{i}$ costs \$$\boldsymbol{A}$ + \$$\boldsymbol{B}$, where \$$\boldsymbol{B}$ represents the potential price premium for the local good.

Individual $\boldsymbol{n}$ derives utility $\boldsymbol{U_{in}}$ by choosing alternative $\boldsymbol{i}$ and $\boldsymbol{U_{jn}}$ by choosing alternative $\boldsymbol{j}$. Formally, consumer utilities $\boldsymbol{U_{in}}$ and $\boldsymbol{U_{jn}}$ can be represented through unobservable indirect utility functions as follows:

$$U_{in} = v(1, I_n - A - B, S_n) + e_{in} \tag{1}$$

$$U_{jn} = v(0, I_n - A, S_n) + e_{jn} \tag{2}$$

where $\boldsymbol{e_{in}}$ and $\boldsymbol{e_{jn}}$ are assumed random components of $\boldsymbol{U_{in}}$ and $\boldsymbol{U_{jn}}$, respectively. $\boldsymbol{S_n}$ represents vector of observable socio-economic attributes of individual $\boldsymbol{n}$ that might affect her/his preferences, and $\boldsymbol{I_n}$ represents income.

To estimate the additional maximum willingness to pay for product $\boldsymbol{i}$, the probability of individual $\boldsymbol{n}$ choosing alternative $\boldsymbol{i}$ is defined as:

$$P_n(i) = Pr(U_{in} \geq U_{jn}). \tag{3}$$

After substituting equations (1) and (2) into (3), we obtain:

$$P_n(i) = Pr\{e_{jn} - e_{in} \leq v(1, I_n - A_n, S_n) - v(0, I_n, S_n)\} \tag{4}$$

Under the assumption that $e_n = e_{jn} - e_{in}$ is logistically distributed, and parameterizing each indirect utility function as a linear function of $S_n$ and $B$, the probability that individual $n$ will choose alternative $i$ can be written as:

$$P_n(i) = [1+exp(\alpha S_n - \delta B)]^{-1}. \tag{5}$$

This is a binary logit model with vector parameters α and scalar parameter δ. For any sample of $N$ individuals, these parameters can be estimated, and using the properties of duality theory, the mean and median price premium for product $i$ can be derived.

## The Survey

During the spring of 2002, five focus groups were conducted across New Hampshire to identify key issues and characteristics of locally produced goods and services. From this information, a survey was designed and pre-tested on 300 individuals at the Made in New Hampshire Expo and around the state. In the summer of 2002, one-thousand surveys were mailed to a representative sample of households across New Hampshire using the series of mailings described in the Dillman Tailored Design Method (Dillman, 2000).[3] The mailings included an announcement letter, followed one week later by a complete survey with personalized cover letter and $1 bill. Households that did not respond to the first survey were mailed a reminder postcard two weeks later, followed by a second survey. After accounting for undeliverables, we received 638 completed surveys, for an overall response rate of 69%. Following the success of the New Hampshire survey, additional funding was obtained, and the study was expanded to Maine and Vermont. During the winter of 2003, two-thousand surveys were mailed to representative samples of Maine and Vermont (1,000 to each state). This resulted in 648 usable

[3] The list of names and addresses for each state was purchased from Survey Sampling, Inc of Fairfield, CT.

surveys from Maine and 660 from Vermont, corresponding to response rates of 70% and 71%, respectively.

The survey began with an identification of the preferences of the respondent towards the state of residence in general, followed by several likert scale questions regarding local and specialty shopping opinions.[4]. Next, respondents were asked if they have purchased various locally produced goods and services in the last 12 months, if they know where to find these items, and if the locations are convenient to them. This was followed by a contingent valuation question that asked about additional willingness to pay for a locally made specialty food item. Half of the surveys referred to a good priced at $5 per unit, while the other half were for a $20 per unit item. The values were chosen through consultation with the staff in the New Hampshire Made program (New Hampshire Stories, Inc) and through a survey of prices of local specialty food products carried in local shops. The question posed to each individual was as follows:

> Let's say you want to buy a specialty food product (maple syrup, salsa, cookies, etc) and saw two kinds in a store. Both were the same quality and cost **$A**. One was made in New Hampshire and one was made out of state. Which would you buy?
>
> __Either one, it doesn't matter __The New Hampshire food __The out of state food
>
> If the food product you chose above cost $ **B** more than the other one, would you still buy it?
>
> ___ YES ___ NO

The bid values for "B" were filled in prior to the survey mailing, and ranged from $1 to $5 for both the $5 and $20 good. The pre-test surveys included a wide range of dollar amounts (from $1 to $20) to find the appropriate range to estimate the willingness to pay for a local product

[4] Respondents were asked to circle a number between 1 (strongly disagree) and 5 (strongly agree), or choose "don't know".

premium. Kanninen (1995) suggests that the dollar amount should fall approximately between the 15th and 85th percentile of "YES" votes, based on pre-testing. After rounding to the nearest dollar, the ranges were approximately equal for the $5 and $20 food items. The last food question asked if the respondent had ever been unhappy with a local specialty food product. The survey finished with a request for socioeconomic information and room for general comments.

## Demographics and Consumer Perceptions

Before describing the consumer perceptions and buying patterns it is useful to know that the local branding programs in Maine, New Hampshire and Vermont are quite different from one another. Maine products are marketed through the “Maine Made: America’s Best” program (see www.mainemade.com), housed in the Maine Department of Economic and Community Development. In New Hampshire, local products and services are marketed through New Hampshire Stories, Inc, a non-profit membership organization, and the “New Hampshire’s Own: A Product of Yankee Pride” slogan (see www.nhmade.com). The Vermont Department of Agriculture, Food and Markets manages the “Vermont Seal of Quality” (see http://www.vermontagriculture.com/aboutsoq.htm).

Table 1 displays a comparison of the respondent demographics with the 2000 Census Bureau Data for the states of Maine, New Hampshire and Vermont. While some statistics are not directly comparable (for example, only adults over 18 were sampled), some differences should be noted. Survey respondents from this study, as in the majority of mail survey research, generally have more education, higher annual income, and are more likely to be male (Miller 1983). This is common for two reasons, but should be noted when extrapolating survey results to the general population. First, when sampling households, one is more likely to address the male head of household in the identification and mailing process. Second, individuals with lower levels of education may have difficulty with the reading and writing of a paper survey.

Table 2 summarizes the results from the section of the survey that questioned knowledge and convenience of locally made agricultural and specialty food products. In order to compare results across states, statistical analysis of the percentage of those that answered "yes" was performed using paired t-tests. Survey respondents from the three states reveal differences in shopping patterns and perceptions of the markets that sell local food products. Of particular note is that the percentage of New Hampshire respondents who know where to purchase state-produced food products or find it convenient to do so is significantly lower than Maine and Vermont in every category. In point of fact, residents of Maine and Vermont tend to be similar to one another at least at the 10% level of significance in every category except knowledge and convenience of specialty food (but not general agricultural) markets. This finding is not surprising given that both Maine and Vermont have relatively well established local good promotion programs that are housed within state agencies and funded by state revenues. On the other hand, New Hampshire's local good promotion program and "New Hampshire's Own" slogan is comparatively new (established in the Fall of 2002) and is not supported by the state but rather a private non-profit organization.

**Willingness to Pay for Locally Produced Specialty Food Products**

In order to estimate the price premium for locally produced specialty food, equation (5) is estimated for a homogeneous \$5 and \$20 specialty food product for each of the three states in northern New England. The binary dependent variable was set to a value of one if and only if the respondent indicated that s/he would purchase the local good with the \$1 - \$5 price premium, and set to zero otherwise. Tables 1 through 3 characterize the raw data used in the analysis.

The vector of socio-economic attributes used in the binary logit model ($S_n$) includes **Prolocal**, a sum of the likert scale questions that indicate the respondent supports buying local goods, the respondent's **Age** in years, median household income (**Inc**), education level in years (**Ed**), number of household members under the age of 18 (**HHyoung**), number of years residing in current state (**Howlong**), a likert scale response to the statement that farmer's markets, a source of specialty food product, are hard to find (**Hardtofind**), and finally the amount of money the local product costs above the non-local food product of equal quality (**Bid2**). Explanatory variables used in the above regression follow the model specified in Loureiro and Hine (2002).

Model results are presented in Table 4 and Table 5 for the two specialty goods. In each case, the proposed local price premium is negatively correlated with willingness to pay for state produced goods, and favorable attitudes towards local goods are positively correlated. Interestingly, the coefficient on education is negative and significant for Maine and Vermont for the \$5 good, but insignificant for the \$20 good and New Hampshire. This contrasts with the model of Loureiro and Hine (2002), who find a positive correlation between education and willingness to pay. However, Govindasamy et al. (1998) and Jekanowski et al.(2000) found that highly educated consumers were the least likely to patronize locally grown produce which lends some support our finding of a negative correlation between education and willingness to pay for state produced goods. The authors offer the following explanations for the negative correlation. First, Govindasamy et al. (1998) believe that the state's labeling and promotion program may have been more popular with young customers and those with less than a high school degree. Jekanowski et al. (2000) find that educated consumers tend to be less susceptible to advertising and branding and hence less receptive to state marketing efforts. Other demographic characteristics are generally insignificant at the 95% confidence level.

Of particular interest is the negative and significant coefficient on the variable indicating that farmer's markets are difficult to find (**Hardtofind**) for the $20 New Hampshire specialty food good. Similar coefficients for the other states with developed local labeling programs are insignificant. One possible explanation is that search costs for New Hampshire consumers are incorporated into the premium value, thus eroding the willingness to pay for the local quality trait. These search costs could presumably be lowered through a promotional campaign designed to inform the average New Hampshire consumer of the location of locally produced specialty goods, including farmers' markets and other venues. This would tend to increase demand for these products.

The key statistics to take away from these models are the willingness to pay estimates. There is no statistically significant price premium identified for New Hampshire and Vermont consumers for the lower-priced $5 good, but a positive premium for Maine. For the more expensive $20 good, all states exhibit significantly positive local price premia, with the point estimate for New Hampshire the lowest of the three. This result indicates that local price premia exist, at least for more expensive specialty food products, and that the use of a state logo has the potential to successfully differentiate state produced specialty food products from imported substitutes.

Moreover, based on the finding of price premia for locally produced specialty goods, an additional policy implication for all of the state labeling programs exists. Brooker and Eastwood (1989) found that just under two-thirds of survey respondents were willing to pay a slightly higher price to cover the labeling expenses costs of the state logo program for tomatoes. Given that consumers in our study are willing to support a price premium to identify state produced specialty foods, the state labeling programs of New Hampshire, Vermont, and Maine may be able to recoup some expenses through increasing prices of state-labeled products. This is a particularly useful finding for the organizers of the New Hampshire's Own program which currently has the lowest level of funding of the three states.

In order to test for differences in WTP between states for identically priced goods and between goods with price differentials, the method of convolution (Poe, Severance-Lossin, and Welsh, 1994) was employed. Table 6 compares the estimated price premia (WTP) for local specialty food products across states, while Table 7 compares the estimated premia for the products with a base price of $5 to the products with a base price of $20 within each state. In general, price premia cannot be distinguished between states for each good, although equivalence between the $5 good for Maine and New Hampshire is rejected at the 95% level of confidence. This suggests that the significantly positive willingness to pay for more expensive locally produced specialty foods does not vary by state, despite disparate marketing and labeling programs for these goods. There are several possible explanations for this result. One is that the sample size for each subregion is such that coefficient estimates are not estimated with the precision necessary to distinguish between the premia. Another is that the promotional programs have more of an effect on the demand for lower priced goods, at least for Maine and New Hampshire (the variance in median WTP for Vermont is unusually large), than for higher valued goods. This could be explained by changes in the elasticity of advertising between lower and higher priced goods.

Furthermore, the results in Table 7 suggest that for New Hampshire and Vermont, willingness to pay for local specialty foods is positively correlated with the base price of the good, or in other words, that the premium is proportional to the base price. Consumers in Maine, however, do not exhibit this pattern, as the price premium for the \$5 good is significantly higher in this state than the others.

## Discussion and Conclusions

This paper uses survey data to examine Northern New Englanders' knowledge and convenience of locally produced specialty food items, and to estimate the willingness to pay for the local quality trait. Maine and Vermont show similarities in buying patterns and perceived convenience of the market locations, while New Hampshire residents show a statistically lower level of purchases and perceived market convenience. A key factor influencing this finding may be that the New Hampshire's state labeling and promotion program is much newer and smaller then those of Maine and Vermont. With more advertising and consumer education, it is expected that over time the differences between New Hampshire, Maine, and Vermont buying patterns and perceived market convenience will become smaller.

Using dichotomous choice contingent valuation methods, we found that consumers of Maine, New Hampshire and Vermont are willing to pay a small price premium for local specialty goods, so long as the price of the good is sufficiently high. However, we were unable to statistically confirm that the price premia differed across States, suggesting that the promotional programs, while different, did not significantly shift demand for these goods. However, there is some evidence that convenient access to local specialty products can affect the premia, most likely through reducing transaction costs.

The research further indicates that residents of Maine display a strong level of loyalty to lower priced food items, with consumers willing to pay a positive premium for the local product.

They are not, however, willing to pay more for the higher priced food item. One potential explanation , following Kahneman and Knetsch's (1992), is that some individuals are willing to pay a fixed dollar amount that fulfils their need for moral satisfaction, and thus the premium will not increase as the base value increases. Unlike their Eastern neighbors, consumers in New Hampshire and Vermont are not found to be willing to pay more for a locally produced, low priced specialty food product, but they will pay a premium for the more expensive good.

As the demand for specialty foods has been especially strong in recent years, state labeling programs have the opportunity to increase profits of local producers if they can effectively promote awareness and loyalty towards these goods. The results of this study should be useful in helping the state labeling and promotion programs of Northern New England understand how specialty goods are perceived by residents and how to promote awareness and loyalty towards these locally produced specialty products. In addition, this paper serves as a demonstration of the contingent valuation method as a tool for deriving consumer willingness to pay measures.

In closing, much research is left to be done with regards to state-labeling programs and processed foods. Possible extensions of this work include identification of the target locally produced specialty good consumer and the characteristics that this group values in the specialty goods they purchase. In addition, it would be interesting to see if New Hampshire residents have changed their preferences since the launch of the "New Hampshire's Own" slogan and labeling system. Resampling New Hampshire residents is scheduled for the summer of 2004.

Table 1: Respondent Demographics

| | Maine | | New Hampshire | | Vermont | |
|---|---|---|---|---|---|---|
| | Actual[a] | Survey | Actual[a] | Survey | Actual[a] | Survey |
| Median Age: | 38.6 | 53 | 37.1 | 53 | 37.7 | 52 |
| Highest Level of Education: | | | | | | |
| Less than 9th Grade: | 5.4% | 4.2% | 3.8% | 2.0% | 5.1% | 2.3% |
| High School Graduate: | 45.5% | 35.5% | 38.8% | 30.0% | 40.8% | 30.2% |
| Associates Degree: | 26.3% | 19.8% | 28.7% | 20.0% | 24.7% | 17.8% |
| Bachelor's Degree: | 14.9% | 22.7% | 18.7% | 28.0% | 18.3% | 27.4% |
| Graduate or Professional Degree: | 7.9% | 17.8% | 10.0% | 20.0% | 11.1% | 22.3% |
| Median Household Income: | $37,240 | $54,958 | $49,467 | $71,606 | $40,856 | $59,687 |
| Gender | | | | | | |
| Male: | 48.7% | 56.9% | 49.2% | 63.1% | 49.0% | 69.2% |
| Female: | 51.3% | 43.1% | 50.8% | 36.9% | 51.0% | 30.8% |
| Avg. Household Size: | 2.39 | 2.6 | 2.53 | 2.7 | 2.44 | 2.5 |
| Children (Under 18) in Household | 0.58 | 0.6 | 0.64 | 0.7 | 0.61 | 0.6 |

[a]US Census Bureau, 2000

Table 2: Comparing Consumer Perceptions of State-Made Food Products across Northern New England

| | Percentage that said "Yes" | | |
|---|---|---|---|
| | Maine | New Hampshire | Vermont |
| Have you purchased a State[a] grown **Agricultural** product in the last 12 months? (fruit, vegetables, dairy, etc) | 94% | 91%[b] | 95% |
| Do you know where to find State[a] grown **Agricultural** products? | 90% | 85%[b] | 93% |
| Do you know where to find State[a] made **Specialty Foods**? | 69%[b] | 52%[b] | 87%[b] |
| Is it convenient to buy State[a] grown **Agricultural** products? | 72%[b] | 67%[b] | 79%[b,c] |
| Is it convenient to buy State[a] made **Specialty Food** products? | 52%[b] | 42%[b] | 72%[b] |
| Have you ever been unhappy with a State[a] **Specialty Food** product? | 15%[b] | 4%[b] | 12%[b,c] |

[a] Surveys listed the specific state name for the respective state.
[b] Statistically different from the other states at the 5% level
[c] Not different from Maine at the 10% level

Table 3: Percentage of Survey Respondents Who Would Buy the Local Food Product

| | Maine | | New Hampshire | | Vermont | |
|---|---|---|---|---|---|---|
| Base Cost of Food | **$5** | **$20** | **$5** | **$20** | **$5** | **$20** |
| Would buy Local Food | 90.9% | 90.6% | 84.6% | 80.6% | 96.2% | 91.3% |
| Would buy Local food if $1 more | 59.4% | 72.4% | 48.3% | 58.1% | 56.8% | 72.2% |
| … if $2 more | 40.0% | 40.0% | 16.7% | 40.0% | 29.2% | 44.0% |
| … if $3 more | 21.2% | 35.5% | 12.5% | 34.8% | 31.2% | 44.1% |
| … if $4 more | 11.8% | 18.2% | 15.2% | 25.8% | 24.2% | 19.4% |
| … if $5 more | 10.3% | 18.2% | 11.4% | 13.8% | 28.1% | 33.3% |

Table 4: Additional Willingness to Pay for a Locally Produced $5 Specialty Food

| | **Maine** | **New Hampshire** | **Vermont** |
|---|---|---|---|
| | **Coefficient (Std Error)** | **Coefficient (Std Error)** | **Coefficient (Std Error)** |
| **C** | -0.297071 | -5.125091 | 0.038359 |
| | (2.554947) | (2.534984) | (1.867771) |
| **Prolocal** | 1.119091 | 1.534976 | 1.112080 |
| | (0.407638) | (0.484037) | (0.360948) |
| **Age** | 0.000318 | 0.033304 | -0.005051 |
| | (0.019495) | (0.020711) | (0.015370) |
| **Inc** | 5.95E-06 | -3.17E-06 | -2.86E-06 |
| | (6.55E-06) | (6.92E-06) | (6.30E-06) |
| **Ed** | -0.234478 | -0.053625 | -0.222000 |
| | (0.126293) | (0.105979) | (0.091513) |
| **Hhyoung** | 0.277211 | 0.175410 | -0.261467 |
| | (0.231538) | (0.254263) | (0.226745) |
| **Howlong** | 0.018809 | -0.011671 | -0.007675 |
| | (0.012293) | (0.015877) | (0.013247) |
| **Hardtofind** | -0.186448 | -0.276840 | -0.072507 |
| | (0.177989) | (0.201927) | (0.161431) |
| **Bid2** | -0.786133 | -0.660883 | -0.287994 |
| | (0.182984) | (0.192657) | (0.144845) |
| LR Statistic | 41.17908 | 29.56798 | 22.72673 |
| LRI | 0.250974 | 0.213386 | 0.134516 |
| Median WTP | $1.46 | $0.13* | -$0.30* |
| 95% C. I. | 0.35 to 2.04 | -3.72 to 1.18 | -19.72 to 1.92 |

*Not significantly different from zero.

Table 5: Additional Willingness to Pay for a Locally Produced $20 Specialty Food

| | Maine | New Hampshire | Vermont |
|---|---|---|---|
| | **Coefficent (Std Error)** | **Coefficent (Std Error)** | **Coefficent (Std Error)** |
| **C** | -4.309647 | 1.528578 | -2.962311 |
| | (2.194235) | (2.482934) | (2.027530) |
| **Prolocal** | 0.849129 | 1.196412 | 1.053427 |
| | (0.431481) | (0.371104) | (0.349444) |
| **Age** | 0.002498 | 0.008662 | -0.000929 |
| | (0.017361) | (0.020171) | (0.018266) |
| **Inc** | 8.01E-06 | 2.02E-06 | 1.02E-05 |
| | (7.19E-06) | (6.90E-06) | (6.40E-06) |
| **Ed** | 0.092303 | -0.182448 | 0.024398 |
| | (0.080633) | (0.108847) | (0.080849) |
| **Hhyoung** | 0.243343 | -0.616967 | 0.047497 |
| | (0.285202) | (0.351017) | (0.270377) |
| **Howlong** | 0.010931 | -0.009862 | -0.032715 |
| | (0.012137) | (0.014113) | (0.014079) |
| **Hardtofind** | -0.004865 | -0.536824 | -0.007514 |
| | (0.158019) | (0.224123) | (0.179898) |
| **Bid2** | -0.635189 | -0.734228 | -0.500996 |
| | (0.160836) | (0.205499) | (0.162948) |
| LR Statistic | 35.56931 | 46.73624 | 38.43466 |
| LRI | 0.198723 | 0.291344 | 0.221172 |
| Median WTP | $1.98 | $1.89 | $2.18 |
| 95% C. I. | 0.95 to 2.62 | 0.53 to 2.50 | 0.53 to 3.02 |

Table 6: Two-Tailed Probabilities
Method of Convolution Test for Equivalence of Median Price Premia Between States

| | Maine | New Hampshire |
|---|---|---|
| $5 Specialty Good | | |
| New Hampshire | 0.0492 | X |
| Vermont | 0.4796 | 0.4917 |
| $20 Specialty Good | | |
| New Hampshire | 0.5159 | X |
| Vermont | 0.4758 | 0.2162 |

Table 7: Two-Tailed Probabilities
Method of Convolution Test for Equivalence of Median Price Premia Between Goods

| | 2-Tailed Significance |
|---|---|
| Maine | 0.2784 |
| New Hampshire | 0.0962 |
| Vermont | 0.0812 |

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# The Effects of Questionnaire Formats on Elicited Preferences and Values in Stated Preference Experiments

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October 11, 2004

The research reported here was supported in part by the Michigan Agricultural Experiment Station and by the Science to Achieve Results (STAR) Program, U. S. Environmental Protection Agency, Grant Number R827922. The authors are solely responsible for any errors. A more complete report of the analysis and results is found in Hoehn, Lupi, and Kaplowitz (2004).

## Abstract

To avoid a net loss of wetland services, Federal and state regulations require mitigation in order to obtain permits for activities that may impair or destroy wetlands. Mitigation raises the policy issue of determining what and how much should be done to offset the loss or impairment of a wetland (NRC 2001). We address this question using a large sample, web-based survey to elicit stated preferences over wetland mitigation projects that provide alternative bundles of ecosystem services. With stated preference methods of survey research, researchers make questionnaire design choices as to how to present complex information to respondents. Psychological research stresses that cognitive constraints lead to characteristic biases when dealing with complicated decisions (Kahneman 2003). One characteristic bias from cognitive complexity is that respondents overweight losses and underweight gains when faced with information that exceeds their ability to understand and assimilate it (McFadden 2000). In this analysis, we examine the effect of two alternative information formats on elicited preferences for wetland mitigation projects. For one treatment group, information on the characteristics of alternative wetland projects is presented to respondents in a tabular format that is typical of many choice experiments. For the other treatment group, the information on wetland characteristics is presented to respondents in a paragraph format similar to that used for many contingent valuation studies. The data are used to estimate pooled and individual random utility models, and tests for a common preference vector are rejected. Consistent with the work of Kahneman (2003) and McFadden (2000), the empirical results from the two information formats revealed that losses in ecosystem quality were over-weighted and gains in quality were under-weighted by respondents using the text-based choice instrument relative to those estimated from the data of respondents using the tabular format. Systematic questionnaire design appears capable of developing choice formats that are robust to characteristic choice biases.

## The Effects of Questionnaire Formats on Elicited Preferences and Values in Stated Preference Experiments

Wetland ecosystems offer an important opportunity for developing ecosystem valuation methods. Wetlands are one of a small number of ecosystem types that are protected and managed under Federal and state regulations. The basic goal of Federal regulation is "no net loss" of wetlands (National Research Council 2001, p. 2). To avoid a net loss of wetland services, Federal and state regulations require mitigation in order to obtain permits for activities that may impair or destroy wetlands. Mitigation raises the policy issue of determining what and how much should be done to offset the loss or impairment of a wetland (Shabman, Scodari, and King, 1996). Typically, losses are offset by actions to restore wetlands in locations near a destroyed or impaired wetland. The degree of mitigation is often based on technical or ecological grounds. However, a purely ecological assessment may not adequately address wetland attributes that are important and valued by human beings. If the latter values are overlooked, a net economic loss may be incurred despite the best "ecological" plans for implementing the no-net-loss goal.

Stated preference methods provide a means of measuring the equivalency of drained and restored wetlands from the viewpoint of economics and human preferences. Stated preference data may be used to estimate the relative values of wetland types and wetland attributes. Such values can enable an analyst to determine how much restored marsh acreage is sufficient to offset the loss of wooded wetland acreage. These values may be summarized in terms of a statistical mitigation equation that gives the wetland acreage to be restored as a function of (a) the acreage of the wetland destroyed and (b) the quality differences between the destroyed and restored wetlands.

Previous research shows that the ecological qualities of wetlands are valued by the general public and that people have different values for different wetland types (Heimlich et al. 1998; Kosz 1996; Phillips, Haney, and Adamowicz 1993; Stevens, Benin, and Larson 1995). Use and nonuse values both have been recognized as important within the overall economic value of wetlands (Woodward and Wui 2001). Previous research is less clear about the values of specific wetland attributes, such as wildlife habitat or access for recreation by the public. Most reported research tends to focus "on the question of 'what *is* the value' and not enough on *what*, in particular, people value" (Swallow et al. 1998, p. 17).

The researchers developed a web-based questionnaire to elicit, through stated preference experiments, the values that the general public associates with common wetlands and wetland attributes. The basic plan was to elicit preferences using choices similar to the mitigation decisions made by regulators in the State of Michigan. The plan was to give respondents pair-wise choices between a wetland that was to be drained and a mitigation plan for restoring a wetland to offset the loss. The question posed to respondents was whether the restored wetland would be enough to offset the loss of the drained wetland. Within the context of the wetland mitigation choice experiment, one of the research objectives was to allow respondents to interact with salient information regarding wetland characteristics and test whether this interaction affects estimated preferences.

A key problem in designing the choice questionnaires was to reduce the perceived complexity of the wetland ecosystem information and choices. Psychological research stresses that cognitive constraints lead to characteristic biases when dealing with complicated decisions (Kahneman 2003). One characteristic bias from cognitive complexity is that respondents overweight losses and underweight gains when faced with information that exceeds their ability to understand and assimilate it (McFadden 2000). In this analysis, we examine the effect of two alternative information formats on elicited preferences for wetland mitigation projects. For one treatment group, information on the characteristics of alternative wetland projects is presented to respondents in a tabular format that is typical of many choice experiments. For the other treatment group, the information on wetland characteristics is presented to respondents in a paragraph format similar to that used for many contingent valuation studies. The data are used to estimate pooled and individual random utility models, and tests for a common preference vector are rejected.

The estimated value results appear consistent with the loss aversion bias common in complex choices (Kahneman 2003; McFadden 2000). Relative to the tabular format, the conventional text format resulting in values that over-weighted losses in ecosystem quality and under-weighted gains. In contrast, the tabular format resulted in in-kind values that were balanced across gains, losses, and quality characteristics.

## Conceptual Framework

Economic values stem from the tradeoffs made by individuals as they attempt to maximize their well-being given their constraints and opportunities (Freeman 1993; Randall 1987). Nonmarket valuation methods, including stated preference methods, provide policymakers with means for accounting for the economic values associated with environmental and natural resource services (Carson et al. 1998; Freeman 1993; Stevens, Benin, and Larson 1995).

By asking individuals to choose among alternative sets of market and nonmarket goods and services, stated preference methods reveal economic values of resource characteristics. Stated preference techniques are widely applied in market research (Louviere et al, 2003), transportation economics (Bates 2000), development economics (Rubey and Lupi 1997), and environmental economics (Adamowicz et al. 1998; Boxall et al. 1996; Mackenzie 1993; Opaluch et al. 1993; Swallow et al. 1998). To be successful, stated preference studies must ensure that respondents understand the environmental goods and services they are asked to value, accept as plausible the proffered policy scenarios, and make economically relevant tradeoffs between alternatives (Arrow et al. 1993; Mitchell and Carson 1989).

Responses to stated preference questions for multidimensional goods such as wetland ecosystems may be analyzed by combining the Lancaster (1971) model of utility with a random utility econometric model (McFadden 1973). The Lancaster model describes utility as a function of goods' characteristics. A good such as a wetland has characteristics that provide service flows to individuals. These service flows affect utility and are valued by individuals. The Lancaster model provides the linkage between the overall value of a wetland and the value of the services derived from its characteristics.

The random utility econometric approach provides a means for statistically linking the Lancaster model to the stated preference rankings for different wetland characteristics. With this approach, data on respondents' preferences for alternative bundles of wetland characteristics may be analyzed using discrete choice methods (McFadden 1973). Individuals' stated preferences for different wetland services reveal the relative utility associated with those services. By varying services, statistical methods are used to estimate the contribution of each service to utility.

In the choice experiments developed by this project, the *i*th respondent, $i \in (1,...,I)$, evaluates a pair of wetlands. Each wetland is described by a vector of wetland characteristics and services, $x_j$, $j \in (1,...,J)$. One of the wetlands described by $x_j$ is to be drained and is referred to as the drained wetland. The second wetland, described by $x_k$, $k \in (1,...,J)$, is a restored wetland offered as mitigation for the drained wetland. The respondent is asked whether the restored wetland makes up for the loss of the drained wetland.

Respondents' choices to the restoration question allow the research to estimate a mitigation equivalency function. The equivalency function is derived in the following way. Each wetland is presumed to yield utility $U_i(x_m)$, $m = j, k \in (1,...,J)$ for the *i*th respondent. Utility is composed of a deterministic function, $u(x_m, c_i)$, of wetland services and individual characteristics, $c_i$, and an unobservable term that is deterministic for the individual, but is unobserved and stochastic for the researcher, $\delta_{im}$,

**(1)** $$U_{im} = u(x_m, c_i) + \delta_{im}$$

While individual utility cannot be observed, restoration choice data provide information on the relative utility of the drained and restored wetlands. The probability that a respondent says that the restored wetland is sufficient to compensate for the loss of the drained wetland is

$$
\begin{aligned}
\textbf{(2)} \qquad Prob(x_k \text{ is chosen}) &= Prob(U_{ik} > U_{ij}, k \text{ ne } j = 1,\ldots,J) \\
&= Prob[u(x_k, c_i) - u(x_j, c_i) > \delta_{ij} - \delta_{ik}, k \text{ ne } j = 1,\ldots,J] \\
&= Prob[\,D(x_k, x_j, c_i) > \varepsilon_{ijk}, k \text{ ne } j = 1,\ldots,J]
\end{aligned}
$$

where $D(x_k, x_j, c_i) = u(x_k, c_i) - u(x_j, c_i)$ is a function representing the difference in the utility between the restored and drained wetlands, and $\varepsilon_{ijk}$ is a stochastic error term. Selecting a functional form for $D(x_k, x_j, c_i)$ identifies the equation and equation parameters for the mitigation equivalency equation. Selecting an empirical distribution for the stochastic error term leads to a particular estimator such as a probit or logit model. The mitigation equivalency equation may be estimated with either the probit or logit formulation using maximum likelihood procedures.

Equation (2) shows how data on mitigation choices may be used to estimate mitigation equivalency equations by solving for $D(\cdot)=0$. These equations summarize respondents' in-kind values for wetland characteristics and qualities. The research developed a web-based questionnaire to obtain the choice data needed to estimate mitigation equivalency equations. Questionnaire prototypes were evaluated and refined in an iterative design process. The iterative process used focus groups, individual interviews, and online pretesting to identify, evaluate, and test alternative questionnaire procedures, formats, and content.

The final questionnaire asked respondents to evaluate five mitigation choices. Each choice was composed of a pair of wetlands. Each wetland was described by its different physical and ecological characteristics. The first wetland was a wetland scheduled to be drained. The second wetland was a wetland that would be restored in order to mitigate the loss of the drained wetland. Respondents were asked whether the restored wetland would make up for the loss of the drained wetland. The web-based questionnaire was used to obtain mitigation choices from a large scale sample of Michigan residents. The resulting choice data were used to estimate mitigation equivalency equations. Each stage of the research is described below.

## Research Procedures

As noted above, wetland ecosystems are complex, with many different attributes that can affect human experience in ways that can be very direct or very subtle. The absence of standardized measures makes it difficult to describe ecosystem change scenarios that are easily understood by the general public. To deal with these issues, the researchers developed an iterative design process described in Figure 1. The first stage of the process addressed the need to understand respondents' baseline knowledge about wetland ecosystems (Hoehn, Lupi, and Kaplowitz 2004).

The second through fourth stages of the design process addressed research Objective 2, the feasibility of a web-based questionnaire. Stage 2 began with the prospective component parts of a questionnaire and used qualitative methods to develop questionnaire prototypes. Stage 3 used qualitative methods to test the prototypes and develop a final questionnaire. Stage 4 adapted the questionnaire to a web-based server. This approach follows the similar approach successfully used in a statewide mail survey of wetland mitigation programs (Lupi, Kaplowitz, and Hoehn 2003; Kaplowitz, Lupi and Hoehn, 2004).

A key step in completing a large sample implementation of the final web-based questionnaire was to obtain an e-mail list of potential respondents. Access to a panel of potential web-based respondents was purchased from Survey Sampling International (SSI) which maintains a panel of potential web-based respondents who have agreed to complete an occasion online questionnaire in return for participation in a prize lottery. The SSI panel is a self-selected sample of potential respondents with known demographic characteristics. Given the types of wetlands addressed by the questionnaire, the sample for the project was drawn from SSI panel members who resided in Michigan.

The web-based survey was implemented in multiple stages beginning in October and ending in December, 2003. The e-mail invitations to the SSI panel resulted in 3,454 clicks on the welcome page to the web-based questionnaire. A large number of individuals chose to leave one or more questions blank during the course of filling out the questionnaire. Usable questionnaires with at least one completed mitigation choice and complete demographic information numbered 1,373 or 40 percent of those visiting the welcome page. Fully complete questionnaires numbered 1,060, or 31 percent of those who clicked on the welcome page. The total number of usable mitigation choices accompanied by usable demographic information was 6,714.

As a point of reference, it may be useful to compare the completion rate on the web-survey with the completion rates for a similar mail survey on wetland mitigation. The mail survey was mailed to 1,500 individuals randomly selected from Michigan Secretary of State's database for licenses and personal identification cards. Using a tailored design (Dillman 2000) with three replacement questionnaires, 62 percent of the mailed questionnaires resulted in either a non-response, a refusal, or a returned questionnaire due to an invalid address. Without excluding the invalid and undeliverable addresses, approximately 35 percent of the mail survey respondents returned questionnaires with complete responses for all five choices (Lupi, Kaplowitz, and Hoehn, 2003).

## Respondents Interaction with Information and Stated Preferences

The web-based questionnaire was designed to facilitate informed wetland mitigation decisions by ordinary respondents drawn from the general public. A key element of the final design was the display of wetland features in a tabular form shown in Figure 2. The tabular form arrayed the relevant wetland choice information in two adjacent columns, one for each wetland under consideration. Wetland habitats were described in four dimensions; habitats for reptiles and amphibians, for song birds, for wading birds, and for wild flowers. Each type of habitat was described with a rating of poor, good, or excellent. The questionnaire contained narrative explaining each of the ratings. The questionnaire provided habitat quality ratings in each of the four habitat dimensions for both the drained and restored wetlands. The ratings were based on what a visitor was likely to see during a visit to the wetland. A "poor" rating meant that the wetland habitat supported "these species in very small numbers...[so] a trained observer is *unlikely to find any* of these species." A "good" rating meant that the wetland habitat supported "these species in average numbers...[so] a casual observer is likely to see a few of these species." An "excellent" rating meant that the wetland habitat supported "these species in better than average numbers...[so] a casual observer is very *likely to see a variety* of these species."

Respondents found the tabular format easy to understand. Pretest comments indicated that the tabular format permitted rapid assimilation of the wetland features, encouraged feature-to-feature comparisons, and facilitated the task of making tradeoffs across different features. A few pretest participants commented that the questionnaire format was "too easy." The working hypothesis at the start of the web-survey was that the tabular format simplified a complex decision in a way that encouraged reasoned, informed decisions. To test this hypothesis, a text version of the questionnaire was developed and included as part of the web-survey implementation plan. This text version of the questionnaire was administered to a randomly selected sub-sample of respondents. The only difference in the text version of the web-based questionnaire was that it replaced the tabular scorecard with two paragraphs of text.

Figure 3 displays the text version of the page comparing the two wetlands. The text version contains information identical to the tabular version. The text version presents wetland features and information is a manner not dissimilar to standard contingent valuation formats. Despite their structural differences, the two versions of the questionnaire contain the same information. The tabular version was successively revised using the iterative design process to simplify the presentation of the information and support respondents' reasoned decisions. The text version did not contain the structural and graphical devices of the tabular version that made it easy for respondent to (a) understand the five different dimensions of habitat quality, (b) attach quality ratings to each wetland and each habitat dimension, and, finally, (c) make tradeoffs across the two wetlands and their habitat qualities.

Previous research suggests the structural differences between the two formats lead to important differences in respondent behavior. Viscusi and Magat (1987) found that tabular formats had a greater impact on risk avoidance behavior and willingness to pay than text formats. Psychological research stresses that cognitive constraints lead to characteristic biases when dealing with complicated decisions (Kahneman 2003). One characteristic bias that has been reported is that respondents overweight losses and underweight gains when faced with information that exceeds their ability to understand and assimilate it (McFadden 2000).

Previous research provided the researchers with a quantitative and testable hypothesis about the performance of the tabular and text questionnaires. If the tabular format was successful in encouraging reasoned, balanced decisions, then respondents should weight wetland gains and losses in a more balanced way, relative to those respondents who used the text version of the questionnaire. In terms of the equivalency functions described in equations (1) to (3), the hypothesis implies that, all else equal, the beta coefficients for gains are larger when estimated on choices for respondents receiving the tabular format than the beta coefficients for gains when estimated on data from the text format. The effect for wetland losses would be just the opposite. The beta coefficients for losses would be smaller for tabular respondents than for respondents who received the text format.

## Tabular and Text Format Results

The tabular and text formats yielded two sets of data suitable for an analysis of mitigation choices and values. The data pertaining to the tabular format was the preferred, core data set, since the tabular design was subject to the full iterative design process. The purpose of the text format data was to provide a baseline for evaluating the effects of the tabular design. By hypothesis, the text format leads to cognitive biases that overweight losses in wetland qualities and underweight gains in wetland quality.

The text and tabular data contained three types of variables. First, there were the wetland choice variables. Respondents were given five mitigation scenarios and were asked to determine whether the restored wetland was sufficient to offset the loss of a drained wetland. Hence, each individual recorded an accept or reject choice for up to five restoration scenarios. Second, there were the variables that described the acreage and qualities of both the drained and restored wetlands. Third, there were demographic variables for each respondent.

Table 1 lists demographic characteristics for respondents to the tabular and text versions of the questionnaire. There were 937 respondents to the tabular version and 363 respondents to the text version who had responses complete enough to the used in the choice analysis. The choice analysis required complete responses for the variables listed in Table 1. Mean levels of income, education, age, and gender were similar for respondents to both the tabular and text versions. One exception was for the age of respondents where the text data set contained about 8 percent more respondents who are over 65 years of age. The mean income level for respondents to both versions was about the same as the 2002 Census mean for the State of Michigan. Respondents to the questionnaires were somewhat more schooled with some college study and were more likely to be female and over 65.[5] Finally, 15 percent of the respondents in each sample had never visited a wetland.

Table 2 lists the wetland characteristics used as variables in each of the wetland choice scenarios. Wetland size was one of the variables and ranged from 5 to 19 acres for the drained wetlands and from 4 to 48 acres for the restored wetlands. The other wetland characteristics were described as categorical variables. The drained and restored wetlands (a) allowed access by the public, denoted by a "yes", (b) allowed access to the pubic with developed trails, denoted by "yes-trails", or (c) made no provision for public access, denoted by "no". The type of wetland was a marsh, a wooded wetland, or a mixture of marsh and woodlands.

[5]The sample selection procedures were intended to be weighted by the Census proportions for males and females in the 2000 Census. However, an error occurred in survey research firm's sample selection process during the waves 1 and 2 of the survey. The error was corrected for waves 3 to 6 and the sample size was increased to meet the demographic criteria for the initial sample design. The increase in sample size was made without cost to the EPA contract. The data for waves 1 and 2 are included in the analysis since gender, as shown below, does not have a significant impact on mitigation choices.

The mitigation choice equation was based on the utility difference approach described in the conceptual section. Let $k$ denote a restored wetland and $j$ denote a drained wetland. For the wetland mitigation choices described by the experimental design, the deterministic function, $D(x_k, x_j, c_i)$, of equation (2) is

**(3)** $$D(x_k, x_j, c_i) = \beta_r a_{ik} + \beta_d a_{ij} + \sum_{g=1}^{10} \beta_g \Delta x_{gi} + \sum_{n=1}^{6} \gamma_n c_{in}$$

where $\beta_r$ is the coefficient of the acreage of the restored wetland, $a_{ir}$; $\beta_d$ is the coefficient of the acreage of the drained wetland, $a_{ik}$; $\beta_g$ is the coefficient of the change in the $g$th wetland characteristic, $\Delta x_{gi}$; and $\gamma_n$ is the coefficient of the $n$th respondent characteristic, $c_{in}$, such as income level or having never visited a wetland. The acreage variables were taken directly from the acreage values appearing in respondents' questionnaires.

The changes in wetland characteristics variables, $\Delta x_{gi}$, were transformations of the data in the questionnaires. The change in access variable indicated whether there was a change in public access in the restored wetland relative to the drained wetland. The change in access variable was given a value of 1 if the restored wetland allowed public access while the drained wetland did not. Change in access was -1 if the restored wetland did not provide for public access while the drained wetland did provide for public access. In other cases, change in access was set to 0.

The change in wetland type variable was a simple, unsigned dummy variable. It was given a value of 1 if there was a change in wetland type between the restored and drained wetlands and set to 0 if there was no change in type. The changes in wetland habitat variables were computed from dummy variables representing the poor and excellent categories. The first step was to assign a dummy variable for each of the poor and excellent quality levels of the drained and restored wetlands. Each of the "poor" dummy variables was given a value of 1 if a particular habitat category was poor in quality, and was set to zero otherwise. Each of the "excellent" dummy variables was given a value of 1 if a particular habitat quality was excellent in quality, and was set to zero otherwise. Dummy variables were created for the "poor" and "good" variables for four habitats (reptiles/amphibians, song birds, wading birds, and wild flowers) and both wetlands, so there were 8 initial dummy variables for quality. The habitat dimension for small animals was kept constant across the choice experiments, so no dummy was created to indicate the quality of habitat for small animals.[6]

The second step in computing the habitat change variables was to compute the difference between the habitat dummy variables between the restored and drained wetlands for each of the four habitat dimensions that varied in the experimental design. For instance, the change in poor dummies for reptiles/amphibians was the difference between (a) the poor reptiles/amphibians dummy for the restored wetland and (b) the poor reptiles/amphibians dummy for the drained wetland. A value of 1 for the latter variable meant that the reptiles/amphibian habitat was poor for the restored wetland and not poor for the drained wetland. A value of -1 meant that the reptiles/amphibian habitat was not poor in the restored wetland and poor in the drained wetland. A value of 0 meant no change in the habitat quality for the reptiles/amphibians habitat across the two wetlands. Similar habitat change variables were computed for computed for the poor and excellent dummies variables for each of the 4 habitat categories, resulting in 4 variables to reflect changes in poor quality habitat and 4 variables to reflect changes in excellent quality habitat.

The respondent characteristics variables, $c_{in}$, were simple levels or categorical dummy variables. Income was measured in thousands of dollars. The remainder of the respondent variables were categorical dummy variables, taking the value of 1 if the respondent had the

[6] The small animals habitat quality was kept constant across the two wetlands to reduce the size of the experimental design. Because the small animals are generalists, this type of habitat was not thought to vary across substantially across the common wetlands under consideration, and the other habitat categories were sufficient to demonstrate the role of habitat quality with respect to respondents' preferences.

characteristic, and taking the value of 0 otherwise. With these definitions, the probability of selecting the restored wetland over the drained wetland is

$$\textbf{(4)} \qquad Prob(\, x_k \; chosen) = Prob[\; \beta_r \, a_{ik} + \beta_d \, a_{ij} + \sum_{g=1}^{10} \beta_g \, \Delta \, x_{gi} + \sum_{n=1}^{6} \gamma_n \, c_{in} > \varepsilon_{ih} \;]$$

An increase in the acreage of the restored wetland was expected to increase the probability of a restored wetland being chosen, so $\beta_r$ was expected to be positive. An increase in the acreage of the drained wetland was expected to decrease the probability of a restored wetland being chosen, so $\beta_d$ was expected to be negative. The coefficients in equation (4) were estimated assuming a normal distribution for the stochastic terms, $\varepsilon_{ih}$. A random effects probit model was used to account for possible correlation between the multiple choices recorded for each individual.

The estimated coefficients for the random effect probit were normalized by dividing each of the coefficients by $\beta_r$, the coefficient of the acreage of the restored wetland. These normalized coefficients may be reinterpreted as the coefficients of a mitigation equation that leaves the mean respondent indifferent between the restored and drained wetland. This mitigation equation is

$$\textbf{(5)} \qquad a_{ik} = - \frac{\beta_d}{\beta_r} a_{ij} - \sum_{g=1}^{10} \frac{\beta_g}{\beta_r} \Delta \, x_{gi} - \sum_{n=1}^{6} \frac{\gamma_n}{\beta_r} c_{in}$$

where $-\beta_d / \beta_r$ is positive since $\beta_d$ is expected to be negative and $\beta_r$ is expected to be positive. Equation (5) provides a means of estimating the amount of restored wetland acreage that compensates the mean respondent given the estimated coefficients and data on the drained wetland acreage, characteristics of the drained and restored wetlands, and mean respondent demographic characteristics. Equations (4) and (5) were estimated for both the text and tabular data. Equation (4) was estimated using the random effects probit estimator. The probit coefficients were then normalized by dividing each by the estimated coefficient for restored wetland acreage. Standard errors for the normalized coefficients were computed using a Wald procedure.

Table 2 displays the estimated normalized coefficients for the mitigation equation (5). The first column in Table 2 lists the names of each of the variables appearing in equation (5). The second and third columns list the estimated normalized coefficients for the tabular and text data. The final column lists the differences between the coefficients of the tabular and text coefficients. The last five rows of Table 2 list the statistical properties of the tabular and text equation estimates.

## Tabular Results

The coefficients for the tabular equation are mostly of the anticipated sign and statistically different from zero at the 95 percent level. Interestingly, the normalized coefficient for drained acreage is not equal to 1, but is 1.42. A coefficient of close to 1 would mean that restored wetland acreage is a very close substitute for drained acreage. However, the coefficient is 42 percent larger than one so the mean respondent requires compensation of 1.42 restored acres for each acre of drained wetland, even when the two wetlands are otherwise presented as equal in access, type, and habitat quality. The premium of 42 percent on the drained wetland acreage is similar to Mullarkey's finding that natural wetlands were more valuable than restored wetlands (Mullarkey 1997). However, Mullarkey found a much larger premium on dollar value of natural wetlands, perhaps due to perceived differences in habitat quality that were not controlled in his work.

Public access and wetland type also have a significant impact on the amount of mitigation acreage that compensate for loss of the drained wetland. The public access coefficient indicates that providing public access reduces the compensating number of mitigated acres by almost 6 acres. A change in wetland type increases the compensating amount of mitigation by almost 5 acres.

The change in habitat variables are all significantly different from zero for the tabular data and have a sign consistent with intuition. Reduction in habitat qualities from good to poor all require additional acreage to offset the loss in quality. A change in reptile/amphibian habitat from good to poor requires more than 8 additional restored acres to offset the loss of quality. A reduction in wild flower habitat from good to poor requires more than 2 acres of additional restored acreage.

Improvements in habitat quality relative to the drained wetland reduce the amount of restored acreage required for mitigation. A change from good wading bird habitat in the drained wetland to excellent habitat in the restored wetland reduces the number of restored acres by more than 5 acres. An improvement from poor habitat in the drained wetland to excellent habitat in the restored wetland is assessed by summing the appropriate coefficients. For instance, for songbirds, a change from poor to excellent reduces the number of restored acres by 6.56 plus 3.80, an overall reduction of 10.36 acres.

Respondent demographic characteristics all tend to affect the level of mitigation that compensates for wetland loss. Increases in respondent income and their schooling tend to reduce the size of compensatory mitigation projects, and having visited a wetland at some point in the past also leads to reductions in the amount of compensating mitigation acres. The latter variable is interesting since it indicates that individuals who have some experience with common wetlands are more inclined to accept the replacement of existing wetlands with restored wetlands.

## Text Results

A notable feature of the statistical results of the text-based equation is the general lack of statistical significance for individual coefficients. The coefficient of drained wetland acreage is almost exactly equal to one, indicating that the text respondents treated restored acreage about equal to drained wetland acreage. However, the coefficient for drained acreage is estimated imprecisely and is not statistically different from zero. The most noticeable feature of the text coefficients is the large size of the poor quality habitat coefficients and the small size of the excellent quality habitat indicators. Respondents who were randomly given the text-based choice question require more acreage compensation for loss in quality than the respondents who were randomly selected to receive the tabular-based choice question. Alternatively, for improvement in restored habitat quality relative to the drained wetland, text respondents behave in just the opposite fashion; they underweight improvements.

The final column of Table 2 shows that these asymmetries are statistically significant for each of the poor habitat coefficients and are significant as a group for the excellent habitat coefficients. Thus, the results suggest that relative to the tabular format the text respondents fell prey to decisions biases that have been noted by psychologists: respondents tend to overweight losses and underweight gains. The tabular questionnaire appears unaffected by such biases. Coefficient estimates are relatively precise and the differences between coefficients seem reasonable and consistent with intuition. The iterative design process appears successful in deriving a questionnaire that supported balanced, reasoned decisions for rather complex mitigation choices.

The strong asymmetry in the resulting data from the text choice questionnaire also appears in estimating mitigation acreage requirements. Suppose one is considering mitigation for the drainage of a 20 acre wetland with good habitat quality in each of the four habitat categories. Consider two restoration projects: the first involves restoration that results in all four habitat qualities being in poor condition, and the second involves restoration that results in all four habitat qualities being in excellent condition. In the first case, the mitigation equation estimated with the tabular data requires 49 acres of restored wetland acres as compensation, but the equation estimated with the text data requires 106 acres of restored wetland as compensation. Conversely, in the second case involving restoration with excellent habitat quality, computing compensating restoration acreage with the tabular equation requires 11 acres of compensation while the text equation requires 28 acres as compensation.

The mitigation examples highlight the differences between the text and tabular data, and the hypothesized superiority of the tabular questionnaire. With the text questionnaire, respondents appear to overweight losses in habitat quality and underweight gains. The underweighting of gains is rather extreme, since the individual habitat coefficients for improvements are small in size and statistically indistinguishable from zero. In contrast, the tabular data results in coefficients that are economically significant, statistically different from zero, much more balanced in their assessment of wetland gains and losses, and accord with the respondent feedback from the focus groups and pretest interviews.

## Conclusions

The research demonstrated that stated choice experiments with complex ecosystems are feasible for the general public. Careful research on baseline knowledge and systematic pretesting appear essential for obtaining reasonable, unbiased stated choice results. The tabular questionnaire format that resulted from a multi-stage, iterative design procedure appeared to perform well. The research also used a simple text-based information treatment as an example of the type of questionnaire that might be developed without the iterative questionnaire design process. The simple text-based questionnaire revealed the kinds of asymmetric biases anticipated on the basis of recent psychological and economic research (McFadden 2000). The more traditional text-based descriptions resulted in losses in ecosystem quality being overweighted and gains in quality being underweighted relative to those estimated using the tabular format. Thus, while ecosystem choices may be complex enough to strain respondents' decision capacities, systematic questionnaire development seem able to help researchers arrive at formats that reduce or eliminate the impact of characteristic biases on the estimated values.

The results demonstrate that wetland qualities and services are valued by members of the general public. From qualitative research, wetland habitats for small animals, birds, and special plants were found to be of special interest and value to respondents (Kaplowitz, Lupi, Hoehn, 2004). Respondents had direct experience with the latter types of wetland habitats and saw them as directly impacted by mitigation activities. The importance of habitat quality emerged consistently at all stages of the research including the initial focus groups, the pilot survey data analysis, the mail survey results, and the web-survey results. This finding is similar to other recent research on wetland ecosystems (Azevedo et al., 2000; Swallow et al., 1998; Stevens et al., 1995).

Two aspects of the research need to be kept in mind in interpreting the results. First, respondents to both the qualitative and qualitative research were drawn from residents of Michigan. Michigan's climate is characteristic of the humid north-central portion of the United States. Wetlands are a common landscape feature, so Michigan residents may have more experience with wetlands than those in other parts of the United States, especially those living in arid regions. Second, while the study provides estimates of how to adjust mitigation ratios to account for differences in habitat quality, it should be considered a first step. The wetlands considered here were common types which are regularly subject to permit actions in Michigan. The study results do not apply to rare wetlands, rare habitats, or rare species. Likewise, in the wetland choices studied here, respondents were explicitly asked to hold other functions of wetlands constant.

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Figure 1. Iterative Web Questionnaire Design and Testing

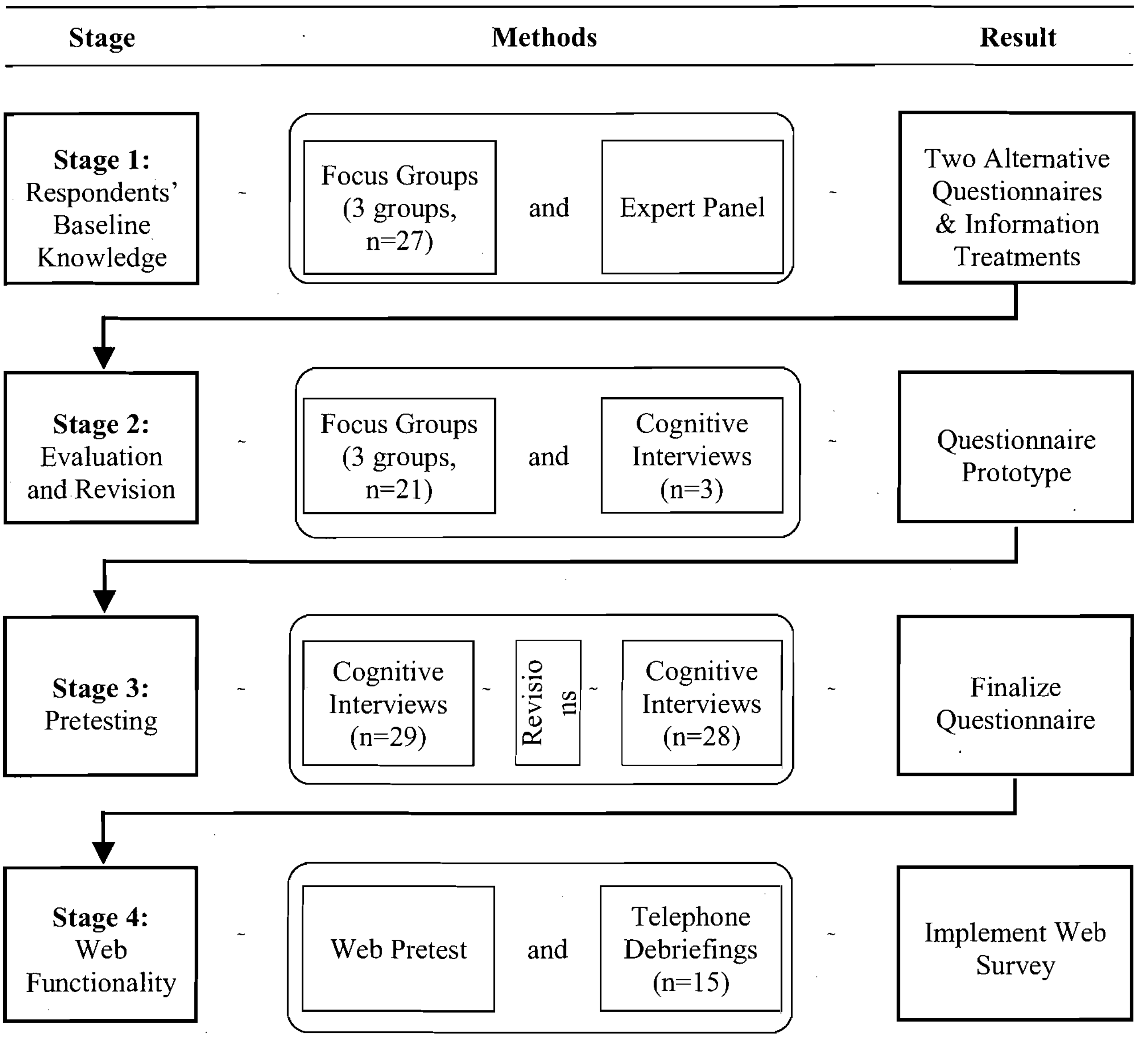

Figure 2. Tabular Choice Format

**Wetlands Scorecard #1**
**How do the Drained and Restored Wetlands Compare?**

| | Wetland Choice #1 | |
|---|---|---|
| **Wetland Features** | **Drained Wetland** | **Restored Wetland** |
| Is it marsh, wooded, or a mix of march and woods? | Wooded | Mixed |
| How large is it? | 14 acres | 23 acres |
| Is it open to public? | Yes | No |
| Are there trails and nature signs? | No | No |
| **How good is the habitat for different species?** | | |
| Amphibians and reptiles like frogs and turtles | Excellent | -- |
| Small animals like raccoon, opossum, and fox | Good | Good |
| Songbirds like warblers, waxing, and vireo | -- | Good |
| Wading birds like sandpiper, heron, or crane | -- | Good |
| Wild flowers? | Good | -- |

Figure 3. Text Choice Format

**Wetlands Scorecard #1**
**How do the Drained and Restored Wetlands Compare?**

**Wetland Choice #1**

**Drained Wetland**

The drained wetland is 14 acres in size. It is a wooded wetland. It is open to the public. It has no trails or nature signs. This wetland is excellent habitat for amphibians. Small animals such as raccoon, opossum, and fox have good habitat in this wetland. The habitat is poor for warblers, waxwing, vireo, and other songbirds. It is poor habitat for wading birds such as cranes, heron, and sandpipers. The growing conditions for wild flowers are good.

**Restored Wetland**

The restored wetland is 23 acres in size. It is a mix of marsh and wooded wetland. It is not open to the public. It has no trails or nature signs. This wetland is poor habitat for amphibians. Small animals such as racoon, opossum, and fox have good habitat in this wetland. The habitat is good for warblers, waxwing, vireo, and other songbirds. It is good habitat for wading birds such as cranes, heron, and sandpipers. The growing conditions for wild flowers are poor.

Table 1. Respondent Characteristics

| Variable | Tabular | Text | Michigan, Census 2000 |
|---|---|---|---|
| Households | 937 | 363 | 3.8 million |
| Income ($1,000) | 54.4 | 54.1 | 57.4 |
| Some college | 79% | 79% | 52% |
| 18 to 25 years | 8% | 8% | 9% |
| Over 65 years | 38% | 47% | 12% |
| Female | 56% | 60% | 49% |
| Never visited a wetland | 15% | 15% | - |

Table 2. Tabular and Text Mitigation Equation Estimates

| Variable | Tabular | Text | Text-Tab |
|---|---|---|---|
| Acreage of drained wetland | 1.42** | 0.99 | (0.42) |
| Change in public access | -5.76** | -9.76** | (3.99) |
| Change in wetland type | 4.69** | 1.81 | (2.86) |
| Change in poor habitat | | | |
| Reptiles/amphibians | 8.19** | 23.46** | 15.3** + |
| Wading birds | 5.76** | 21.11** | 15.3** + |
| Song birds | 6.56** | 21.33** | 14.8** + |
| Wild flowers | 2.33** | 12.51** | 10.2** + |
| Change in excellent habitat | | | |
| Reptiles/amphibians | -4.76** | 1.00 | 5.8+ |
| Wading birds | -5.09** | (1.12) | 4.0+ |
| Song birds | -3.80** | (1.76) | 2.0+ |
| Wild flowers | -1.94** | 3.44 | -.5+ |
| Income ($1,000s) | -0.06** | (0.03) | 0.30 |
| Some college | -4.25** | 3.91 | 8.20 |
| 18 to 25 years of age | 2.53 | 3.35 | 0.80 |
| 65 years of age and over | 0.41 | (3.29) | (3.70) |
| Female | -2.9* | 0.59 | 3.50 |
| Never visited a wetland | 8.26** | (0.54) | (8.80) |
| Intercept | 4.75 | 6.68 | 1.90 |
| Group variance to total variance | 0.37** | 0.39** | - |
| No of observations | 4685 | 1811 | - |
| % correct yes | 63 | 64 | - |
| % correct no | 64 | 65 | - |
| Log-likelihood | -2814 | -1151 | - |

A "*" indicates that a coefficient is significantly different from zero at 90% level; "**" means significantly different from zero at 95% level; and a "+" that a coefficient is significantly different from zero when evaluated as a group of habitat variables.

# Modeling Preference Asymmetries in Stated Preference Data: An Application to Rural Land Preservation

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March 1, 2004

*Presented at the 2004 Annual W-1133 Meetings, February 25-28, Maui, HI.*

This research was funded by the USDA Fund for Rural America 97-36200-5219. Opinions belong solely to the author and do not imply endorsement by the funding agencies or the U.S. Government. Address correspondence to the author at the Connecticut Sea Grant College Program, University of Connecticut at Avery Point, 1080 Shennecossett Road, Groton, CT 06340-6048 or rjohnston@canr.uconn.edu.

**Modeling Preference Asymmetries in Stated Preference Data: An Application to Rural Land Preservation**

## Abstract

In their simplest form, response asymmetries imply that different factor weightings determine the extent to which respondents support versus oppose otherwise identical statements. This paper assesses whether response asymmetries occur in common Likert scale ratings of policy preference and implications of such asymmetries for estimating determinants of heterogeneity in policy support. The empirical application assesses determinants of land use policy preferences among Rhode Island rural residents. We present a model that may be applied when preference asymmetries are present and contrast this alternative to traditional ordered response models which assume response symmetry. Results indicate that assumptions of response symmetry implied by standard models may prevent detection of significant influences on policy preference.

## Introduction

Public preferences for land use or other public policies are often elicited using variants of the common Likert scale (e.g., Bateman et al. 2002; Danielson et al. 1995; Kline and Wichelns 1998; Lynne et al. 1988; Variyam et al. 1990). Such scales are designed to provide information regarding a respondent's strength of preference above and beyond a simple binary (e.g., yes or no) response. However, in return for the ability to model the increased information provided by Likert scales (LS), researchers often accept implicit assumptions not required when modeling binary responses. These include the assumption that respondents choose a cardinal response on the continuum of the provided scale by reference to a single underlying preference function.

Despite the common use of simplifying assumptions when working with LS data, literature addressing other choice contexts suggests that responses to such preference scales may be somewhat more complex. For example, as Likert scales often allow respondents to express varying degrees of either support or opposition, there is the possibility that responses will manifest preference or response asymmetries (Yamagishi and Miyamoto 1996; Shafir 1993). Response asymmetries formally imply that different factor weightings determine the extent to which respondents support versus oppose otherwise identical statements, and might also cause the determinants of an initial binary choice (e.g., oppose versus support) to differ from determinants of preference intensity (e.g., how strongly do I oppose or support).

The only formal discussion of response asymmetries in the stated preference literature is provided by Johnston and Swallow (1999), who show that asymmetries may occur in more complex, two-stage stated preference questions. An example of a two-stage stated preference question would be one in which respondents are first asked whether they support or oppose a hypothetical policy, then asked for their strength of support or opposition. In this context, Johnston and Swallow (1999) demonstrate that different functions may govern the extent to which respondents support versus oppose hypothetical watershed management plans—an extension of prior experimental findings reported in the psychology literature (Yamagishi and Miyamoto 1996; Shafir 1993). However, while Johnston and Swallow (1999) demonstrate the existence of response asymmetries in two-stage questions, they fail to provide a practical modeling alternative that allows for such behavioral patterns. Hence, as admitted by the authors, the analytical and policy guidance provided by their empirical results is somewhat limited.

This paper assesses whether similar behavioral patterns manifest in (much more common) stated preference questions designed to be answered in a single stage; a classic example is the LS rating. Specifically, we assess implications of preference asymmetry for cases in which statistical models are used to assess heterogeneity in stated preferences associated with demographic or other attributes of individual respondents. For example, one might wish to assess heterogeneity in support for particular land use policy tools associated with attributes such as age, income, and education. These models are typically estimated using ordered logit or probit, with the LS response as an independent variable (cf., Swallow et al. 2001). A discovery of response asymmetries in such common choice frameworks would represent a significant and perhaps surprising finding, as it would imply that even simple ordered preference ratings along a single, continuous preference scale (e.g., a Likert scale) may involve more complex choice processes than are currently anticipated or modeled in the literature.

Simply put, when estimating ordered response models of LS data, one typically assumes that responses are symmetric. In empirical terms, ordered response models presume that the weight given to independent variables—as revealed by estimated coefficients—is approximately constant over the range of possible outcomes, subject to increasing or decreasing returns and/or interactions captured by the chosen functional form. A corollary to this assumption is that respondents choose a LS rating with a single-stage process. However, other choice mechanics are possible. For example, when presented with a LS question regarding support for a land use policy, respondents might (without prompting) first assess whether they support or oppose the policy, then assess the strength of their support or opposition. In such cases, LS responses may no longer be symmetric, and typical ordered response models may provide improper inferences regarding the impact of independent variables on policy support or opposition.

In cases where response asymmetries are evident, alternative choice models may reveal behavioral patterns obscured by traditional modeling approaches. To address such possibilities, this paper presents a model of LS responses that may be applied when preference asymmetries are suspected, and contrasts this alternative to a more traditional ordered response approach. This allows both formal hypothesis tests for the presence of preference asymmetries and assessments of policy implications.

## A Standard Strength of Preference Model

Our application concerns estimation of the relationship between attributes of survey respondents and stated preferences for common land use policy tools. Preference for each tool is measured on a standard LS, in which respondents are asked to rate each policy tool on a five-point scale ranging from "strongly oppose" (1) to "strongly support" (5).

Standard random utility models assume that a respondent's strength of preference for a given policy tool (or statement) is determined by the utility that would result from the application of that tool, compared to the utility generated by the status quo, or lack of that tool. Here, the attributes of each specified policy tool are fixed. Hence, we suppress the characteristics of each tool within the utility function; utility is specified as a sole function of characteristics of the individual or household. That is, for each management tool *i*, the difference in utility resulting from the application of that tool may be specified

$$dU_i = dv_i(\mathbf{D}) + \theta_i \tag{1}$$

where $dv_i(\mathbf{D})$ is the deterministic or observable component of the utility difference, **D** is a vector of variables characterizing demographic and other characteristics of the individual or household hypothesized to influence management preferences, and $\theta_i$ is the stochastic element of the utility difference. Equation (1) models heterogeneity in preferences for specific policy tools, as a function of individual and household attributes.

Ordered response models represent a standard approach to such problems. This approach presumes that the individual assesses the utility difference $dU_i$ associated with a particular policy tool, then indicates within which of a set of intervals this utility difference falls. Here, each interval corresponds to a specific LS response represented by a strength of preference indicator variable $L_{ij}$, where $j = \{1, 2, ..., 5\}$, such that

$$
\begin{aligned}
L_{ij} &= 1 \text{ if } \alpha_{i,j-1} < dU \leq \alpha_{i,j} \qquad (2)\\
&= 0 \text{ otherwise.}
\end{aligned}
$$

For example, if the respondent "strongly opposes" a management tool, then $L_{i1} = 1$, and $L_{i2} = ... = L_{i5} = 0$. The $\alpha_{ij}$ in (2) represent utility thresholds associated with particular values of $L_{ij}$. These thresholds are unobserved and treated as parameters by the ordered response model (Maddala 1983).

Following Johnston and Swallow (1999), equations (1) and (2) allow one to estimate the probability that a respondent provides a particular strength of preference response (i.e., that the difference in utility is in category $j$):

$$\Pr(\alpha_{i,j-1} < dU_i \leq \alpha_{i,j}) = \Pr(dU_i \leq \alpha_{i,j}) - \Pr(dU_i \leq \alpha_{i,j-1}) \text{ for } j=1, 2, \ldots, 5. \quad (3)$$

$$= [1 - \Pr(dv_i - \alpha_{i,j-1} \leq \theta_i)] - [1 - \Pr(dv_i - \alpha_{i,j} \leq \theta_i)]$$

$$= \Pr(\theta_i < dv_i - \alpha_{i,j-1}) - \Pr(\theta_i < dv_i - \alpha_{i,j})$$

where Pr is the probability operator. Assumptions regarding the distribution of $\theta$ determine whether (3) is estimated as an ordered probit or ordered logit model, with appropriate likelihood functions provided by Maddala (1983), among others. Here, we estimate the model using an ordered probit likelihood function. We emphasize the fact that such models estimate parameters defining a single preference function, $dv_i$, applicable to responses over the entire continuum represented by the Likert scale.

## A Model Allowing for Preference Asymmetries

In contrast to the standard approach exemplified by (1)-(3) above, models incorporating preference asymmetry would allow judgments of *how strongly do you support* to be determined by a different choice mechanism or component weighting than judgments of *how strongly do you oppose*, following a binary choice in which an option is either supported or opposed (Yamagishi 1996; Yamagishi and Miyamoto 1996). For example, when faced with the opportunity to rate a policy tool on a continuous LS, a respondent might first make an initial (binary) decision to support or oppose the tool in question. Simultaneously, the respondent would choose a level of support or opposition. These two choices, however, need not be governed by an identical choice or preference functions.

The model specifies one preference or choice function to govern the initial support versus oppose choice,

$$dU_{1i} = dv_{1i}(\mathbf{D}) + \theta_{1i} \tag{4}$$

where the subscript '1' denotes the support versus oppose choice. The respondent's choice is represented by the indicator variable $L_j$ which here takes on a value of 0 if the respondent opposes policy tool $i$ and 1 if the respondent supports the tool. We denote this as the "first-stage" choice, although it may be made simultaneously to the "second-stage" choice revealing preference intensity.

The second-stage choice reveals a respondent's preference intensity, or strength of support or opposition. We consider that it is made simultaneously to the first-stage choice, although one might also consider it a subsequent choice. The preference intensity choice is assumed to be governed by the function

$$dU_{2i} = dv_{oi}(\mathbf{D}) + dv_{si}(\mathbf{D}) + \theta_{2i} \tag{5}$$

where the subscripts $o$ and $s$ represent 'oppose' and 'support', $dv_{oi}(\mathbf{D}) = 0$ for $L_i = 1$, and $dv_{si}(\mathbf{D}) = 0$ for $L_i=0$. Simply put, for those whose first-stage choice indicates support for the policy tool in question, $dU_{2i} = dv_{si}(\mathbf{D}) + \theta_{2i}$. For those whose first-stage choice indicates opposition, $dU_{2i} = dv_{oi}(\mathbf{D}) + \theta_{2i}$.

We assume that strength of support or opposition is revealed through a binary choice, represented by the indicator variable $S_i$. Respondents who support tool $i$ may choose to "moderately" ($S_i$=0) or "strongly" ($S_i$=1) support. Respondents who oppose tool $i$ may choose to "strongly" ($S_i$=0) or "moderately" ($S_i$=1) oppose. Note that $S_i$ is defined for support and oppose contexts such that the directional effect of utility difference on strength of preference is preserved.[1]

The special case, implied by the standard ordered response model (1)-(3), is that $dv_{1i}(\mathbf{D}) = dv_{oi}(\mathbf{D}) = dv_{si}(\mathbf{D})$—allowing the two-component decision governed by (4) and (5) to be collapsed into a single-component decision governed by (1). However, while it is certainly possible that the same underlying preference or utility function determines support or opposition, strength of support (for supported tools), and strength of opposition (for opposed tools), a preference asymmetry model considers this an hypothesis subject to testing. That is, it would also allow for the case in which $dv_{1i}(\mathbf{D}) \neq dv_{oi}(\mathbf{D}) \neq dv_{si}(\mathbf{D})$, as suggested by (4)-(5) above.

As a result of the additional flexibility characterized by (4)-(5) above, models incorporating preference asymmetries allow for a possibilities not often considered by the stated preference literature. For example, standard ordered response models assume that each demographic indicator (e.g., age) has a fixed marginal effect on utility difference function, and that this single function determines the LS rating of the entire continuum. Implicit in this approach is the assumption that both the magnitude and directional impact (i.e., sign) of each indicator is fixed.

Such behavioral assumptions notwithstanding, there are a variety of other systematic mechanisms through which demographic and other indicators may systematically influence stated preferences. Perhaps the most obvious is that certain demographic attributes may be associated with stronger (or more moderate) preferences for any given policy choice, regardless of whether that policy choice is supported or opposed. For example, older respondents may tend to have more extreme opinions than younger residents—stating stronger opposition to disliked policies and stronger support for favored policies, *ceteris paribus*.

Were such patterns to hold, the marginal directional effect of age on strength of preference (i.e., the sign of the coefficient) would change as one moved from the "oppose" to "support" segment of the LS continuum. Because standard ordered response models of LS data do not allow for this possibility, misspecification of respondents' choice behavior is possible. A typical symptom of such misspecification would be that the model would fail to identify a systematic effect of a particular attribute (e.g., age) on strength of preference, when in fact a systematic and significant effect exists.

*The Econometric Model*

While estimation of ordered response models for LS data is well established, appropriate modeling of preference asymmetries in such data has (to the authors' knowledge) little precedent in the literature.[2] Moreover, the model characterized by (4)-(5) above lends itself to a variety of existing estimation methods, depending on the behavioral assumptions and data manipulations that one is prepared to accept.[3]

Here, we model the choices implicit in (4) and (5) as simultaneous bivariate decisions with correlated disturbances, in the tradition of seemingly unrelated regressions, where correlation is incorporated by $\rho=\text{Cov}[\theta_{1i}, \theta_{2i} \mid \mathbf{D}]$ (Greene 2003). The first bivariate choice, corresponding to (4), indicates a respondent's opposition ($L_i=0$) or support ($L_j=1$) for a specific management tool $i$. The second bivariate choice, corresponding to (5), indicates a respondent's strength of preference

($S_i$={0,1}), where statistical determinants of this choice may differ between those opposing and supporting tool $i$ as noted above.

Non-independence between the two choices may be incorporated by assuming a bivariate normal distribution of equation errors, leading to estimation using a bivariate probit likelihood function (Poe et al. 1997). Assuming a linear specification of $dv_{1i}$ in (4),

$$dU_{1i} = dv_{1i}(\mathbf{D}) + \theta_{1i} = \beta_{1i}\mathbf{D} + \theta_{1i}\,. \tag{6}$$

Here, $\beta_{1i}$ is a conforming vector of coefficients to be estimated. We assume a similar linear specification of (5), such that

$$dU_{2i} = dv_{oi}(\mathbf{D}) + dv_{si}(\mathbf{D}) + \theta_{2i} = \gamma_{oi}\mathbf{D_o} + \gamma_{si}\mathbf{D_s} + \theta_{2i} \tag{7}$$

where $\mathbf{D_o}$ is a vector of individual specific attributes for respondents who oppose tool $i$ (i.e., $\mathbf{D_o}$=0 for $L_i$ =1), $\mathbf{D_s}$ is a vector of individual specific attributes for respondents who support tool $i$ (i.e., $\mathbf{D_s}$=0 for $L_i$ =0), and $\gamma_{oi}$ and $\gamma_{si}$ are conforming vectors of coefficients to be estimated.

Assuming that $\theta_{1i}$ and $\theta_{2i}$ are distributed $N(0,\ \sigma_1^2)$ and $N(0,\ \sigma_2^2)$, respectively, and defining $z_1 = \beta_{1i}\mathbf{D}/\sigma_1$ and $z_2 = (\gamma_{oi}\mathbf{D} + \gamma_{si}\mathbf{D})/\sigma_2$ as standardized normal errors (cf. Poe et al. 1997), the standard bivariate normal distribution for $z_1$ and $z_2$ is given by

$$\Phi(z_1, z_2, \rho) = \frac{e^{-(z_1^2 + z_2^2 - 2\rho z_1 z_2)/2(1-\rho^2)}}{2\pi(1-\rho^2)^{1/2}}\,. \tag{8}$$

Given (8), parameters are estimated using a readily estimable bivariate probit model; the likelihood function for this model is provided by Greene (2003) and Poe et al. (1997), among others.

Model estimation allows for readily accessible hypothesis tests of various aspects of potential preference asymmetry. For example, the test of null hypothesis $H_0$: $\gamma_{oi} = \gamma_{si}$ assesses whether the statistical determinants of strength of opposition to tool *i* are significantly different from those of strength of support. Moreover, comparisons of the overall fit and performance of the ordered response and bivariate probit models allow appraisals of each model's ability to appropriately characterize respondents' choice behavior.

*Treatment of Neutral Responses*

Although the bivariate probit approach provides significant flexibility in allowing for preference asymmetries, it does so at a cost. Specifically, as the model is specified as a combination of two bivariate decisions, it cannot incorporate neutral responses, described in the survey as "neither oppose nor support". Recall, the data of interest is comprised of LS ratings on a five point scale, where the median score (3) represents a neutral response to tool *i*. The ordered response model (1) – (3) is able to incorporate such data points within model estimation, as part of the continuum of LS responses. However, the bivariate probit model models a binary "oppose versus support" choice jointly with a binary strength of preference choice; it does not incorporate neutral responses, which are dropped from the data prior to estimation. Hence, certain information (data) is lost in estimating the bivariate probit model.[4] This is not unique to this approach to preference asymmetries; Johnston and Swallow (1999) also drop neutral responses in hypothesis tests of preference asymmetries in two-stage stated preference questions. Nonetheless, it is important to view the performance of the bivariate probit model in light of the smaller dataset from which it is estimated. In theory, the additional information incorporated in the ordered response models should afford additional efficiency and robustness. However, such advantages may be offset if such models misspecify respondents' behavior.

## The Data

Data are drawn from the *Rhode Island Rural Land Use* survey, an instrument designed to assess rural residents' tradeoffs among attributes of residential development and conservation. Survey development required over eighteen months, including background research; interviews with policy makers and residents; and focus groups (Johnston et al. 2002). Surveys were mailed to 4000 randomly selected residents of four Rhode Island rural communities following the total survey design method (Dillman 2000). Of 3702 deliverable surveys, 2157 were returned, for a response rate of 58.2%. Further details of the survey and its administration are provided by Johnston et al. (2002; 2003).

Survey respondents were asked to indicate their degree of opposition to, or support for twenty-one different land use management policy options, on a five-point LS ranging from 'strongly oppose' (1) to 'strongly support' (5). Policy options included zoning changes, fee-based land preservation techniques, tax policies, housing caps, impact fees, and other land use policy tools common in Rhode Island rural communities. Based on the results of focus groups, all policies were described in simple, non-technical terms. Table 1 lists the policy options rated by respondents, and the mean support ratings associated with each option. Mean scores above 3.0 indicate that the average respondent supports the policy option, with higher scores indicating greater mean support. Mean scores below 3.0 indicate that the average respondent opposes the policy option, with lower scores indicating greater mean opposition. Diversity in average responses across similar management tools suggests that respondents considered each policy in detail when providing LS responses, rather than providing identical ratings of broadly similar policies (e.g., tools 1 and 2; tools 7-9).

## Empirical Results

Empirical models compare performance of the ordered probit (traditional) and bivariate probit (preference asymmetry) approaches, applied to the same LS data. For both models, responses are modeled as a function of an identical set of independent variables. Independent variables include length of residency in the rural community, standard demographic descriptors characterizing age, income, and education, and other indicators such as membership in environmental or business organizations or ownership of a local home (table 2). As the attributes of each rated management tool are fixed (table 1), they are not incorporated in the statistical models.

Distinct ordered and bivariate probit models are estimated for each of the 21 management tools considered by respondents, resulting in a total of 42 estimated models. Table 3 summarizes overall model statistics including the likelihood ratio $\chi^2$ for each model, McFadden's pseudo-$R^2$ for both models (McFadden 1974), and the likelihood ratio $\chi^2$ for the null hypothesis $H_0$: $\gamma_{oi} = \gamma_{si}$ (i.e., that the statistical determinants of strength of support are identical to those for strength of opposition, in the bivariate probit model).[5]

All models are statistically significant at better than $p<0.01$, as indicated by likelihood ratio tests (table 3). Differences in the data used in the bivariate and ordered probit estimations (i.e., elimination of neutral responses from the bivariate model) preclude standard specification tests and direct comparisons of log likelihood functions. Nonetheless, model fit statistics provide support for the bivariate probit model. For example, both overall model $\chi^2$ and pseudo-$R^2$ statistics are improved in the bivariate probit specifications. Indeed, the average pseudo- $R^2$ increases by more than a factor of three compared to the ordered probit model.

Bivariate probit results also show strong evidence of preference asymmetry in strength of preference responses. The null hypothesis $H_0$: $\gamma_{oi} = \gamma_{si}$ (i.e., that strength of preference for supported tools is determined by an identical attribute weighting as strength of preference for opposed tools) is rejected at $p<0.01$ in all cases (table 3). Hence, we see strong evidence that determinants of strength of preference for opposed tools differs from analogous determinants for supported tools—violating one of the primary assumptions upon which traditional ordered response modeling relies. The presence of such asymmetries may help explain the relatively poorer statistical performance of ordered probit relative to the bivariate probit in this context. Hence, while direct specification tests are infeasible, and despite the larger dataset from which the ordered response model is estimated, the general fit of the bivariate probit model of LS responses appears to improve over that of the traditional (i.e., ordered response) approach.

*Implications for Heterogeneity in Policy Preferences*

Additional insight regarding the policy relevance of such results may be gained by reviewing model results for specific management tools. Given the impracticality of illustrating full estimation results for each of the 42 estimated models, we focus discussion on a small number of cases. Although we emphasize cases in which evidence of preference asymmetry is relatively clear, similar evidence may be found in most estimated models. This evidence suggests that asymmetric responses may have considerable and meaningful impacts on the results of standard ordered preference models.

For example, table 4 illustrates results for tool 1 and tool 6, including both bivariate probit and ordered probit models. As noted in table 4, tool 1 is described as "attract new commercial development to your town by offering tax incentives." For tool 1, both the bivariate and ordered models are statistically significant at better than $p<0.001$, based on likelihood ratio tests (table 4). Results of the ordered probit model suggest that the following four attributes influence strength of preference (or utility) at $p<0.10$: length of residency (positive influence); age (positive); membership in an environmental organization (negative); and membership in a business group (positive).

In contrast, the bivariate probit model allows attribute influence to differ depending on the choice being made. For the support versus oppose choice, where larger estimates of $dv_{1i}(\mathbf{D})$ are associated with a greater probability of supporting the management tool in question, the bivariate model finds statistically significant influence associated with six attributes, including: length of residency (positive); age (positive); gender (female respondents associated with more negative responses); home ownership (positive); membership in an environmental organization (negative); and membership in a business group (positive).

The results of both models are intuitive. For example, members of environmental organizations might be expected to state greater opposition to tax incentives designed to attract commercial development, while members of business groups might be expected to express greater support. However, the bivariate model is able to discern statistically significant effects (at least in the support versus oppose model) for two additional attributes: gender ($p<0.01$) and home ownership ($p<0.05$). Based on bivariate probit results, these attributes influence the probability of supporting commercial tax incentives. While the ordered probit p-values for these attributes are close to the generally accepted $p=0.10$ threshold for statistical significance, we nonetheless cannot reject the individual null hypotheses of zero influence on LS responses (table 4).

Policy implications of such results are not difficult to envision. For example, if a policymaker were to request information on support for commercial tax incentives among local homeowners, the traditional approach to LS data (ordered response modeling) would indicate no statistically significant influence—a result of potential importance when seeking to identify constituencies for particular policy options. However, the bivariate strength of preference model suggests that this conclusion may be misleading. Based on bivariate probit results, one would conclude that homeowners are more likely to support such tax incentives at $p<0.05$.

The potential rationale for this discrepancy is straightforward. The bivariate model indicates that home ownership influences both the probability of supporting commercial tax incentives (*own_home*, table 4), as well as strength of opposition for those respondents who oppose such policies (*own_home×oppose*; $p<0.05$). However, home ownership cannot be shown to influence the strength of support among those who support such policies (*own_home×support*; $p=0.21$). The lack of a statistically significant effect over a *portion* of the LS continuum likely contributes to the failure of the ordered probit model to identify a statistically significant effect of home ownership over the entire response continuum.[6]

Aside from an improved ability to identify statistically significant attribute effects, the bivariate model also reveals differences in preference determinants among those who oppose and those who support tool 1 (table 4). For example, significant effects on strength of opposition ($dv_{oi}(\mathbf{D})$) are associated with residence duration (positive or *weaker* opposition), female respondents (negative or *stronger* opposition), age (negative), home owners (positive), and members of environmental groups (negative). In contrast, strength of support ($dv_{si}(\mathbf{D})$) is associated with age (positive or *stronger* support), female respondents (negative or *weaker* support), members of environmental groups (negative), and members of business organizations (positive). Hence, as suggested by the joint hypothesis test in table 3, the bivariate probit model for tool 1 indicates that determinants of strength of support and strength of opposition differ.

Bivariate strength of preference results for tool 1 also reveal a characteristic incidence of preference asymmetry associated with the variable *age*. As noted above, older residents who oppose tool 1 reveal stronger opposition at $p<0.01$ (*age×oppose*<0; table 4). However, older residents who support tool 1 reveal stronger support at $p<0.03$ (*age×support*>0). Combining these results leads to the conclusion that increasing age is associated with stronger preferences for commercial tax incentives—both in support and opposition—a classic representation of response asymmetry associated with a demographic attribute.

Similar results are evident for tool 6 (table 4). To streamline discussion of these results, table 5 provides a simplified illustration of statistically significant effects identified by each model, with 'plus' and 'minus' signs indicating positive and negative statistically significant impacts. As shown by table 5, the signs of statistically significant effects in the bivariate support/oppose model are identical to those found in the ordered probit model—a sign that both models are identifying similar patterns in LS responses. However, among various symptoms of response asymmetry manifest in the bivariate strength of preference model for tool 6, table 5 provides another archetypal illustration of preference asymmetry and its potential implications.

The identified preference asymmetry is associated with the attribute *house_size* (the number of people in the household). Household size cannot be shown to influence the probability of supporting the "revitalization of town centers using public funds" (tool 6). However, for those respondents who support tool 6, larger household sizes are associated with stronger support at

$p<0.01$ (tables 4,5). In contrast, for those respondents who oppose tool 6, larger household sizes are associated with stronger opposition at $p<0.08$. Such patterns cannot be captured by standard ordered response specifications—despite the statistically significant patterns identified by the bivariate model, the ordered probit model shows *house_size* to have an insignificant effect on LS responses.

Here again, we find a pattern of potential relevance obscured by the ordered response framework: members of larger household tend to express stronger preferences regarding revitalization of town centers. If one opposes such policies he/she will oppose more strongly; if one supports he/she will support more strongly. Aside from indicating patterns of heterogeneity in policy support, these results also have potential implications for the implicit weight given to respondents from larger households in analysis of LS responses. That is, the tendency of such respondents to provide more extreme (or outlier) responses may provide them with a greater-than-average influence on statistical results.

A final illustration of response asymmetries is provided for tool 8, described in the survey as "purchase and preserve undeveloped land with public funds." A summary of statistically significant effects is provided by table 6. Again, for simplicity we emphasize only the direction (sign) of statistically significant effects.[7] Here, the ordered probit model identifies preference heterogeneity associated with only three out of nine attributes: *age*, *hi_educate*, and *envi_group*. In contrast, the bivariate probit model—including both the support/oppose and strength of preference models—identifies statistically significant impacts with six out of nine attributes: *age*, *female*, *house_size*, *hi_educate*, *envi_group*, and *bus_group*. Here, the ability to distinguish attribute effects on the support/oppose choice versus the strength of preference choice allows the identification of additional sources of response heterogeneity, in this case associated with household size, gender, and membership in business organizations. For example, results indicate an negative effect of household size on the probability of supporting the purchase and preservation of undeveloped land ($p<0.05$); the statistical significance of this effect is not apparent in the ordered response model. For policymakers or researchers interested in forecasting referendum support for proposed policies among different demographic groups, such patterns may be of considerable relevance.

Similar patterns are found (to varying degrees) in models addressing LS responses for all 21 management tools considered. Results strongly support the hypothesis that response asymmetries occur, thereby refuting a primary assumption upon which standard ordered response models rely. However, perhaps more importantly, results show that alternative choice models—here a bivariate probit specification—are able to identify behavioral patterns obscured by traditional ordered response models of LS data. These findings illuminate aspects of preference heterogeneity that—while of questionable policy relevance in some cases—may be of considerable importance in particular policy or analysis contexts.

## Conclusion

If one is solely interested in calculating mean policy support over a sample of respondents, then findings of preference asymmetry in LS responses may be of little relevance. However, if one wishes to characterize heterogeneity in support for management tools or assess statistical determinants of LS responses, then the potential for such patterns may be of critical importance. Here, we show that the assumption of response symmetry implied by common ordered response models may prevent detection of potentially significant influences on strength of support or opposition, as revealed by LS data. Model results support the potentially surprising conclusion that statistically significant response asymmetries are both common and policy relevant, even in relatively straightforward LS ratings applied over a single ordered preference scale.

Aside from establishing the general existence of response asymmetries in our LS data, findings here indicate that preferences for land use tools are potentially more complex than is typically assumed. For example, bivariate probit specifications of our LS data frequently reveal differences among those variables influencing the decision to support or oppose particular land use policies and those variables influencing strength of preferences for supported or opposed policies. Moreover, while certain universal and intuitive patterns are apparent (e.g., environmental group membership is almost universally associated with stronger support for pro-environment policies and stronger opposition to pro-development policies), the effect of other attributes varies considerably. Such patterns suggest caution in making general statements concerning heterogeneity in preferences for particular types of land use policies.

Although the present analysis provides evidence that response asymmetries occur in simple LS questions (i.e., questions designed to be answered in a single stage, on a five-point continuum) and discusses potential policy implications of such patterns, there is much left for future research. For example, researchers often use principal component factor analysis of the response correlation matrix to estimate latent factors that capture a high degree of variation in LS responses. Resulting factor scores are then used as either independent or dependent variables in statistical models (e.g., Variyam et al. 1990). Implications of response asymmetry (in the raw LS data) on derived factor scores—and on statistical models incorporating these scores—has yet to be explored.

Additional areas of future research include alternative approaches to data incorporating response asymmetries. Bivariate probit models represent only one potential means to model response asymmetries of the type identified here. Other potential approaches might be drawn from variants of Heckman-type sample selection models (e.g., Greene 2003) or nested choice models (e.g., nested logit). While exploration of these and other potential approaches to LS data is beyond the scope of this paper, we emphasize that the bivariate model estimated here is only one potential approach. Other, as-yet-undeveloped approaches might (or might not) provide superior means to model LS response asymmetries of the type identified here (e.g., alternative approaches might allow preference asymmetries to be modeled while allowing retention of neutral responses). The potential performance of potential alternative approaches notwithstanding, results here suggest that models allowing for response asymmetries in LS data may provide considerable insight over and above that provided by traditional ordered response models.

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**Table 1. Likert Scale Strength-of-Support Ratings for Land Use Policy Options[a]**

| Option | Description (survey text) | Mean Rating (Std. Dev.) |
|---|---|---|
| 1 | Attract new commercial development to your town by offering tax incentives | 2.49 (1.26) |
| 2 | Attract new residential development to your town by offering tax incentives | 1.85 (0.94) |
| 3 | Encourage preservation by reducing property taxes on undeveloped land | 4.11 (0.89) |
| 4 | Encourage new development by expending public water and sewer services | 2.31 (1.11) |
| 5 | Discourage people from moving into your town by increasing the tax rate | 1.94 (0.89) |
| 6 | Revitalize town or village centers using new public funds | 3.36 (1.03) |
| 7 | Purchase and preserve undeveloped land with private funds (e.g., land trust donations) | 4.08 (0.81) |
| 8 | Purchase and preserve undeveloped land with public funds (e.g., public bond issues) | 3.58 (1.05) |
| 9 | Purchase and preserve undeveloped land through a new real estate sales tax | 2.68 (1.16) |
| 10 | Collect fees from developers to offset costs of additional public services for new developments | 4.16 (0.86) |
| 11 | Collect fees from developers to offset additional environmental damages from new developments | 4.28 (0.83) |
| 12 | Encourage residential development by decreasing zoning restrictions | 1.79 (0.92) |
| 13 | Encourage commercial development by decreasing zoning restrictions | 1.95 (1,04) |
| 14 | Require new developments to preserve some undeveloped land | 4.21 (0.76) |
| 15 | Require trees and shrubs between new houses and roads | 4.11 (0.82) |
| 16 | Further protect water resources by increasing zoning restrictions | 4.08 (0.83) |
| 17 | Further protect wildlife resources by increasing zoning restrictions | 4.04 (0.87) |
| 18 | Require new commercial development to occur along major roadways | 3.75 (1.01) |
| 19 | Require new commercial development to occur within town or village centers | 3.00 (1.09) |
| 20 | Institute a cap on the total number of new homes allowed to be built each year | 4.09 (0.93) |
| 21 | Tighten enforcement of existing zoning and subdivision regulations | 4.02 (0.86) |

[a] Measured on a five-point Likert-scale in which 1 = "strongly oppose" and 5 = "strongly support." Numbers in parentheses are standard deviations.

**Table 2. Variables Included in the Strength of Preference Models**

| ***Variable Name*** | Description | Units and Measurement | Mean (Std. Dev.) |
|---|---|---|---|
| *Resid_year* | Length of residency in the community in which a respondent currently resides. | Number of Years | 16.59 (15.55) |
| *Age* | Reported age of survey respondent. | Number of Years | 47.28 (12.44) |
| *Female* | Dummy variable distinguishing male and female respondents. | Binary (0,1), 1 = female; 0 = male. | 0.33 (0.47) |
| *House_size* | Size of household, including children. | Number of individuals. | 2.93 (1.30) |
| *Own_home* | Dummy variable identifying those who indicate that they own their principle residence (versus renting). | Binary (0,1), 1 = home owner. | 0.91 (0.28) |
| *Hi_educate* | Dummy variable identifying those respondents with at least a four-year college education. | Binary (0,1), 1 = four-year college or greater education. | 0.34 (0.47) |
| *Hi_income* | Dummy variable identifying those respondents with reported household income above $39,999 per year. | Binary (0,1), 1 = income > $39,999. | 0.53 (0.49) |
| *Envi_group* | Dummy variable identifying those indicating membership in environmental groups (Audubon Society, land trusts, etc.). | Binary (0,1), 1 = environmental group member. | 0.19 (0.39) |
| *Bus_group* | Dummy variable identifying those indicating membership in business organizations (chambers of commerce, etc.). | Binary (0,1), 1 = business group member. | 0.20 (0.40) |

**Table 3. Model Statistics: Ordered Probit and Bivariate Probit Estimation Results**

| Model (Policy Tool) | Ordered Probit LR $\chi^2$ (df=9)[a] | Ordered Probit Pseudo-$R^2$ | Bivariate Probit LR $\chi^2$ (df=28)[a,b] | Bivariate Probit Pseudo-$R^2$ | LR $\chi^2$ for $H_0$: $\gamma_{oi} = \gamma_{si}$ (p-value) |
|---|---|---|---|---|---|
| 1 | 93.64 | 0.017 | 192.48 | 0.047 | 76.00 (0.01) |
| 2 | 96.00 | 0.021 | 111.96 | 0.035 | 38.58 (0.01) |
| 3 | 63.65 | 0.014 | 163.30 | 0.048 | 103.71 (0.01) |
| 4 | 101.70 | 0.019 | 162.96 | 0.045 | 70.96 (0.01) |
| 5 | 26.06 | 0.006 | 79.00 | 0.027 | 43.00 (0.01) |
| 6 | 48.77 | 0.009 | 173.99 | 0.054 | 118.26 (0.01) |
| 7 | 97.79 | 0.023 | 172.10 | 0.055 | 100.92 (0.01) |
| 8 | 100.20 | 0.019 | 219.37 | 0.062 | 117.45 (0.01) |
| 9 | 94.13 | 0.016 | 206.82 | 0.055 | 89.95 (0.01) |
| 10 | 39.93 | 0.009 | 119.96 | 0.037 | 61.83 (0.01) |
| 11 | 43.41 | 0.011 | 115.92 | 0.035 | 54.21 (0.01) |
| 12 | 111.55 | 0.025 | 110.16 | 0.034 | 29.18 (0.01) |
| 13 | 115.73 | 0.024 | 153.25 | 0.043 | 28.82 (0.01) |
| 14 | 55.87 | 0.014 | 135.56 | 0.043 | 63.04 (0.01) |
| 15 | 69.84 | 0.016 | 150.39 | 0.051 | 88.80 (0.01) |
| 16 | 44.85 | 0.010 | 140.67 | 0.046 | 93.87 (0.01) |
| 17 | 75.30 | 0.016 | 153.47 | 0.049 | 92.21 (0.01) |
| 18 | 48.92 | 0.010 | 135.45 | 0.040 | 74.79 (0.01) |
| 19 | 51.67 | 0.009 | 169.25 | 0.047 | 96.63 (0.01) |
| 20 | 37.82 | 0.008 | 115.89 | 0.034 | 67.27 (0.01) |
| 21 | 57.71 | 0.013 | 146.52 | 0.049 | 92.67 (0.01) |

[a] All models are statistically significant at $p<0.01$.

[b] Statistics are for the full bivariate probit model including both equations.

**Table 4. Ordered and Bivariate Probit Results: Tools #1 and #6[a]**

| | Bivariate Probit: Support/Oppose | | | | | |
|---|---|---|---|---|---|---|
| | Tool #1 | | | Tool #6 | | |
| Variable Name | Parameter Estimate | Std. Error | p>\|z\| | Parameter Estimate | Std. Error | p>\|z\| |
| *Resid_year* | 0.012 | 0.003 | 0.001 | 0.005 | 0.003 | 0.045 |
| *Age* | 0.005 | 0.003 | 0.067 | -0.012 | 0.003 | 0.001 |
| *Female* | -0.189 | 0.072 | 0.008 | 0.272 | 0.079 | 0.001 |
| *House_size* | -0.023 | 0.026 | 0.387 | 0.010 | 0.030 | 0.742 |
| *Own_home* | 0.266 | 0.130 | 0.041 | -0.316 | 0.140 | 0.024 |
| *Hi_educate* | -0.067 | 0.074 | 0.364 | -0.078 | 0.079 | 0.330 |
| *Hi_income* | 0.001 | 0.073 | 0.992 | -0.001 | 0.080 | 0.990 |
| *Envi_group* | -0.260 | 0.091 | 0.004 | 0.139 | 0.098 | 0.156 |
| *Bus_group* | 0.228 | 0.082 | 0.005 | 0.158 | 0.093 | 0.088 |
| *Intercept* | -0.763 | 0.186 | 0.001 | 0.874 | 0.218 | 0.001 |
| | Bivariate Probit: Strength of Preference[b] | | | | | |
| *Resid_year × Oppose* | 0.009 | 0.002 | 0.001 | 0.009 | 0.004 | 0.015 |
| *Age × Oppose* | -0.008 | -0.002 | 0.001 | -0.017 | 0.005 | 0.001 |
| *Female × Oppose* | -0.137 | 0.068 | 0.045 | 0.178 | 0.117 | 0.129 |
| *House_size × Oppose* | -0.016 | 0.021 | 0.448 | -0.074 | 0.042 | 0.076 |
| *Own_home × Oppose* | 0.242 | 0.121 | 0.045 | -0.322 | 0.222 | 0.148 |
| *Hi_educate × Oppose* | -0.087 | 0.072 | 0.225 | -0.138 | 0.108 | 0.203 |
| *Hi_income × Oppose* | -0.091 | 0.071 | 0.196 | -0.018 | 0.113 | 0.873 |
| *Envi_group × Oppose* | -0.249 | 0.085 | 0.004 | 0.158 | 0.138 | 0.254 |
| *Bus_group × Oppose* | 0.093 | 0.082 | 0.259 | 0.023 | 0.125 | 0.855 |
| *Intercept × Oppose* | 0.853 | 0.110 | 0.001 | 2.157 | 0.306 | 0.001 |
| *Resid_year × Support* | 0.003 | 0.003 | 0.422 | 0.005 | 0.003 | 0.076 |
| *Age × Support* | 0.011 | 0.005 | 0.023 | -0.005 | 0.004 | 0.267 |
| *Female × Support* | -0.215 | 0.120 | 0.074 | 0.105 | 0.092 | 0.252 |
| *House_size × Support* | -0.004 | 0.047 | 0.933 | 0.097 | 0.034 | 0.005 |
| *Own_home × Support* | 0.293 | 0.233 | 0.209 | -0.303 | 0.147 | 0.040 |
| *Hi_educate × Support* | -0.009 | 0.114 | 0.940 | 0.056 | 0.096 | 0.558 |

| | | | | | | |
|---|---|---|---|---|---|---|
| *Hi_income* $\times$ *Support* | 0.011 | 0.120 | 0.929 | 0.067 | 0.097 | 0.489 |
| *Envi_group* $\times$ *Support* | -0.404 | 0.166 | 0.015 | 0.081 | 0.113 | 0.476 |
| *Bus_group* $\times$ *Support* | 0.333 | 0.119 | 0.005 | 0.178 | 0.105 | 0.091 |
| *Intercept* $\times$ *Support* | -2.108 | 0.345 | 0.001 | -1.456 | 0.267 | 0.001 |
| $\rho$ | 0.999 | 0.001 | 0.001 | 0.998 | 0.332 | 0.004 |
| *N* | 1648 | | | 1401 | | |
| -2 LnL $\chi^2$ (df=28) | 192.48 | | 0.001 | 173.99 | | 0.001 |
| | | | Ordered Probit | | | |
| *Resid_year* | 0.009 | 0.002 | 0.001 | 0.006 | 0.002 | 0.001 |
| *Age* | 0.007 | 0.002 | 0.003 | -0.007 | 0.002 | 0.003 |
| *Female* | -0.083 | 0.053 | 0.119 | 0.193 | 0.053 | 0.001 |
| *House_size* | 0.024 | 0.021 | 0.243 | 0.026 | 0.020 | 0.207 |
| *Own_home* | 0.142 | 0.092 | 0.123 | -0.249 | 0.091 | 0.006 |
| *Hi_educate* | -0.065 | 0.056 | 0.246 | -0.007 | 0.055 | 0.894 |
| *Hi_income* | -0.005 | 0.055 | 0.928 | 0.014 | 0.055 | 0.794 |
| *Envi_group* | -0.224 | 0.066 | 0.001 | 0.049 | 0.064 | 0.446 |
| *Bus_group* | 0.163 | 0.064 | 0.011 | 0.127 | 0.064 | 0.047 |
| Estimated Cut-Points | | | | | | |
| $\alpha_1$ | -0.112 | | | -1.602 | | |
| $\alpha_2$ | 0.640 | | | -0.904 | | |
| $\alpha_3$ | 1.085 | | | -0.076 | | |
| $\alpha_4$ | 2.175 | | | 1.281 | | |
| *N* | 1886 | | | 1899 | | |
| -2 LnL $\chi^2$ (df=9) | 93.64 | | 0.001 | 48.77 | | 0.001 |

[a] The text describing tool #1 is: "attract new commercial development to your town by offering tax incentives." Tool #6 is described as "revitalize town or village centers using new public funds."

[b] For bivariate probit strength of preference model oppose responses, 0=strongly oppose and 1=moderately oppose. For support responses, 0=moderately support and 1=strongly support (see text for additional information)

**Table 5. Summary of Statistical Results: Tool 6[a]**

| Variable Name | Bivariate Probit: Support / Oppose | Bivariate Probit: Strength of Preference Support | Bivariate Probit: Strength of Preference Oppose | Ordered Probit |
|---|---|---|---|---|
| *Resid_year* | + | + | + | + |
| *Age* | - | | - | - |
| *Female* | + | | | + |
| *House_size* | | + | - | |
| *Own_home* | - | - | | - |
| *Hi_educate* | | | | |
| *Hi_income* | | | | |
| *Envi_group* | | | | |
| *Bus_group* | + | + | | + |

[a] A '+' indicates a statistically significant positive effect. A '-' indicates a statistically significant negative effect.

**Table 6. Summary of Statistical Results: Tool 8[a]**

| Variable Name | Bivariate Probit: Support / Oppose | Bivariate Probit: Strength of Preference Support | Bivariate Probit: Strength of Preference Oppose | Ordered Probit |
|---|---|---|---|---|
| *Resid_year* | | | | |
| *Age* | - | | - | - |
| *Female* | | + | | |
| *House_size* | - | | | |
| *Own_home* | | | | |
| *Hi_educate* | + | + | | + |
| *Hi_income* | | | | |
| *Envi_group* | + | + | + | + |
| *Bus_group* | | | - | |

[a] Tool 8 described as "purchase and preserve undeveloped land with public funds (e.g., public bond issues)." A '+' indicates a statistically significant positive effect. A '-' indicates a statistically significant negative effect.

## Endnotes

[1] That is, for both oppose and support contexts $S_{ij} = 1$ corresponds to a higher level of utility or preference $dU_{2i}$ within each category. One may also envision this intuitively as splitting the data (observations) into support (S) and oppose (O) responses, creating two independent datasets of binary strength of preference responses. The support (S) data include all "support" and "strongly support" responses (LS responses 4, 5); the oppose (O) data include all "strongly oppose" and "oppose" responses (LS responses 1,2). The resulting support and oppose datasets are then vertically "stacked", or pooled into a single binary dataset, such that the directional effect of the utility difference on strength of preference is preserved. The result is a pooled dataset incorporating both the oppose and support data vertically stacked. For examples of similar data transformations, see Mazzotta and Opaluch (1995) and Johnston and Swallow (1999).

[1] Johnston and Swallow (1999) present hypothesis tests that establish the presence of preference asymmetries in two-stage stated preference questions, yet fail to provide a consistent approach that would allow one to model respondents' choices in the presence of such asymmetries.

[1] Perhaps the most straightforward approach to this issue would be to apply generalized ordered logit, an approach which relaxes the proportional odds assumption implicit in traditional ordered logit models (cf. US EPA 2002). We eschew this model in favor of an alternative specification (bivariate probit) which formalizes the hypothesized two-stage decisions implicit in a preference asymmetry model. We thank Scott Shonkwiler for suggesting this alternative approach.

[1] As a practical matter, neutral responses make up a very small proportion of the Likert scale data in question. However, some data is nonetheless discarded in estimating the bivariate probit model.

[1] The $\chi^2$ statistic is calculated as $-2[LR_R - LR_U]$, where $LR_R$ is the log likelihood function of the restricted model in which $\gamma_{oi} = \gamma_{si}$, and $LR_U$ is the log likelihood function of the unrestricted model (7).

[1] Recall, the ordered probit model estimates only one parameter estimate per attribute, which applies over the entire continuum of Likert scale responses.

[1] Full results for all models are available from the authors upon request.

# Do Nearby Forest Fires Cause a Revision in Residential Property Values? Results of a Hedonic Property Value Analysis

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March 19, 2004

Acknowledgements: I would like to thank Adam Orem for assistance in assembling the property datasets. Comments from Regional Research Project W1133 members at our annual meeting improved this paper. This research was supported by USDA Forest Service, Pacific Northwest Research Station under Research Joint Venture PNW-03-JV-11261955-266, and the Colorado Agricultural Experiment Station under Regional Research Project W1133.

**Do Nearby Forest Fires Cause a Revision in Residential Property Values? Results of a Hedonic Property Value Analysis**

## Abstract

We used a hedonic property regression model to determine if there was a statistically significant decrease in property values in the town of Pine, Colorado, that was two miles from a major wildfire that burned 12,000 acres, and destroyed 10 houses in the small town of Buffalo Creek. Using a hedonic property model specification similar to that used to evaluate the effect of earthquakes or floods on house prices, we find the forest fire in the Buffalo Creek area had a statistically significant negative effect on house prices in the nearby unburned community of Pine. The house price drop is about 15% using both linear and semi-log hedonic regression specifications. This suggests that home buyers and sellers appear to upwardly revise their perceptions of fire risk after a major fire, reducing the desirability of living in the forest, and reducing house prices relative to before the fire.

Keywords: forest fires, hedonic property analysis, non-market valuation

## Introduction and Problem Statement

An important question for setting public policy is whether people update their risk knowledge of low probability events with the occurrence of natural events such as fires and floods. That is, do these events result in an information feedback by which homeowners reduce their demand for houses in high hazard natural areas? Given government emergency aid and homeowners insurance, perhaps homeowners are financially protected from such events, and only by eliminating emergency aid can homeowners be made to face the full cost of their housing location decisions. However, perhaps there are uninsured wealth losses or utility reductions from increased risk to life from these events. There may also be losses in utility from the harshness of the post natural disaster landscape (e.g., post-fire or post-flood). If so, then after these events the reduced desirability of living in what has now been revealed to be a more hazardous location should be reflected in the housing market.

There have only been a few studies comparing pre- and post natural-disaster event property values. The natural hazards studied have included earthquakes (Brookshire, et al. 1985; Murdoch, et al. 1993) and floods (Damianos and Shabman, 1976; Shultz and Fridgen, (2002)) but have not included forest fires. Forest fires are becoming a more regular multi-billion dollar natural hazard due to increased housing construction in the wildland urban interface zone. Knowing if housing markets incorporate updated risk information is important for determining what, if any, changes in government policy are necessary. If housing prices in areas adjacent to recent forest fires decrease, then markets may be efficient at signaling the increased forest fire risk, reducing the need for changes in government policy in the wildland urban interface zone.

The Hedonic Property Method (HPM) is an appropriate tool for investigating whether there have been statistically significant changes in property values and the magnitude of that change. Because the natural experiment is before and after a forest fire, we use the hedonic price function to control for any changes in house characteristics and other exogenous trends during this time period. Similar to Murdoch, et al.'s approach for earthquakes and Shultz and Fridgen (2002) approach for floods, the effect of the forest fire is tested using a pre-post fire dummy variable. Like Shultz and Fridgen we also test for coefficient equality in the pre and post fire period using a Chow-Test for hedonic property model coefficient equality pre-post fire.

## Hedonic Property Method

### *Theory*

The hedonic property model begins with a consumer who derives satisfaction or utility from housing and all other goods. The utility from housing is a function of the structural characteristics of the house (S1...Sm), non-environmental characteristics of the general neighborhood such as school quality, demographic composition (N1...Nn) and location specific environmental amenities such as forests, wildlife habitat, etc. (Z1...Zi). Consumers are assumed to maximize their overall utility subject to their budget constraint, which is defined over their income and market prices.

In the housing market, the neighborhood and environmental attributes are largely fixed by the location and are inherent in the parcel itself. These attributes cannot be repackaged *per se*, and therefore buyers compete among themselves for parcels that have higher levels of the desirable characteristics. This bids up the price of land parcels that have these desirable features (such as lakeshore properties in Langsford and Jones (1995) analysis), yielding insights regarding the private willingness to pay (WTP) for these characteristics.

If the housing market is competitive such that buyers and sellers decisions cannot affect the equilibrium price in the market area, then following Freeman (1993), the hedonic price function takes the general form:

(1) $Ph= f( S1..Sm; N1..Nn; Z1..Zi)$

Regressing the property price, Ph, against these attributes yields coefficients on each variable. These regression coefficients measure the rate of change in the property price with respect to a one unit change in each attribute, that is, $\partial Ph/\partial Z_i$ , and are the implicit marginal prices of the attribute. These can been used to infer households' WTP for marginal changes in the level of parcel attributes.

In our analysis we are interested in whether the marginal implicit price for a forested property changes when a wildfire occurs far enough away not to threaten the structure, but close enough that it makes quite clear that such a fire could affect their property and the forest around it. As formalized below, we will test this by whether the hedonic price function in (1) shifts down after the wildfire in a town outside of the immediate fire area, but close enough that there may be a good information transfer.

*Empirical Specification*

To isolate the effect of the wildfire on property values, it is important to control for other non-forest related amenities of the property such as the structural attributes. Thus our empirical specification of the hedonic price function is:

(2) Real House Price =

Bo+B1(Acreage) +B2(Fireplace) + B3(Garage) +B4(#Bathrooms) +B5(YearBuilt) +B6(Trend) +B7(PostFire)

Where:

Acreage is the size of the land parcel the house is located on

Fireplace= a dummy variable for whether the house had a fireplace

Garage= a dummy variable for whether the house had a garage

#Bathrooms= the number of bathrooms

YearBuilt= the year in which the home is built

Trend = continuous date variable to reflect underlying trend in house prices in the study area

PostFire= whether the property sold prior or after (=1) the forest fire

The dependent variable is the sale price in constant dollars (1983=100) using the housing price index portion of the Consumer Price Index. Note we do not include any neighborhood attributes because we are studying just one community and they do not vary across properties or over our short time period of analysis (eight years).

The rationale and sign of most of the variables is fairly intuitive, as larger acreages provide higher levels of forest amenities, wildlife habitat, recreation opportunities as well as privacy. A fireplace, garage and additional bathrooms are desirable structural features.
The YearBuilt variable controls for age of the house, and we hypothesize a positive sign, reflecting the desirability of newer homes, as well as the fact that homes depreciate overtime. As suggested by Murdoch, et al., a continuous measure of sale date to reflect the underlying real estate trend is included to control for other exogenous factors during this time period.

The PostFire variable is one for houses selling at least 60 days after the fire event. We chose 60 days as the actual decision to purchase the home is typically made two months prior to the actual sale date being recorded due to a 60 day escrow period.

The simplest functional form for the hedonic price function is linear. With this functional form the marginal implicit price of the characteristic is simply its coefficient. Thus the linear model has easily interpreted and transparent marginal prices. However, the linear form has some drawbacks of constant marginal implicit prices and assumes the consumer can repackage characteristics.

Non-linear functional forms for the hedonic price function overcome these limitations and provide marginal implicit prices for a characteristic that depends on the level of that particular characteristic and on the level of other characteristics as well. Potential non-linear models include the semi-log transformation of the dependent variable and more generalized Box-Cox transformations. The Box-Cox transformation makes the interpretation of the marginal values less intuitive as the attributes are raised to exponents and it makes calculation of the marginal values far more cumbersome (Lansford and Jones, 1995: 343). Cropper, et al. performed a simulation exercise evaluating the accuracy of different functional forms against a known true function. They found that simpler functional forms such as linear and semi log transformation outperformed more complex functional forms in the face of omitted variable bias or use of proxy variables in place of theoretically correct variables (Cropper, et al. 1988).

It is likely that our empirical application shares some of the features mentioned by Cropper, et al. that makes simpler functional forms desirable. Specifically, due to multicolinearity among some of the housing characteristics, we are able to include only a subset of these, and hence the included ones act as proxies for related measures of housing attributes (e.g., bedrooms and overall house size are omitted due to high correlation with number of bathrooms). Based on the arguments of Cropper, et al. we adopt a semi-log model for our non-linear functional form but retain the linear to provide a more directly interpretable measure of marginal willingness to pay from the regression coefficients as well as test the sensitivity of results to different functional forms. As shown below our results are not sensitive to choice of linear or semi-log functional form.

## Hypothesis Tests

Effect of Fire on House Prices--To test whether there is a downward revision in property values after a nearby wildfire, we will evaluate the statistical significance and magnitude of the coefficient on PostFire coefficient, B7. The statistical significance will be evaluated using a standard t-statistic. In the linear model, the magnitude of the drop in property value is given by the coefficient itself, B7. In the semi-log model the magnitude of the drop in property value is given by B7* Real House Price (Taylor, 2003: 354).

Effect of Fire on the Entire Hedonic Price Equation--To test whether one or more of the hedonic price function slope coefficients change after the fire, we test for coefficient equality of the Pre and Post hedonic price function. This coefficient equality test is conducted using a Chow-test, which is an F-test based on comparing the sum of the residual sum of squares of the equations estimated separately for each time period and a single equation estimated using the data pooled across both time periods.

## Study Area

Data for this analysis comes from sales of houses in the town of Pine, which is about two miles from the May 1996 Buffalo Creek fire. The Buffalo Creek fire started on the nearby Pike-San Isabel National Forest and burned parts of the town of Buffalo Creek in Colorado. The Buffalo Creek fire was one of the first major forest fires in the 1990's, after decades of successful fire suppression. The high intensity wildfire burned 12,000 acres, and destroyed 10 houses in the small town of Buffalo Creek. The loss of living forest vegetation and the mineralization of the soil from the high fire temperatures led to a destructive flash flood two months later that closed the main highway and destroyed the town's water treatment system.

## Data

To assess whether the Buffalo Creek fire caused a reduction in house prices in the nearby, but unburned town of Pine, we collected house sales data for three years prior to the fire (1993-1996), and five years after the fire from Jefferson County, Colorado. The data were error checked to verify that transactions were actual arms length transactions. This resulted in over 500 observations.

## Statistical Results

Table 1 presents the linear and log dependent variable hedonic price functions. In both regression, the adjusted R square is 50% or slightly higher, and all the housing variables had intuitive signs and were significant at the 1% level. The pre-post fire variable was negative and statistically significant at the 5% level in both regressions. The linear model loss estimate is $17,095 per house (in constant 1980 dollars) or 15% of the house price. The semi-log model estimated a similar loss of $18,519, representing 16% of the house price.

The percentage reduction in property value due to increased likelihood of wildfire is in the range of estimates that past studies have found for earthquakes (up to –10% in Murdoch, et al. 1993) and –10% for flood events (Shultz and Fridgen, 2002). It may be that recovery of the forest from fire is a longer process than fixing structural damage from earthquakes or floods.

Related to this is recovery of the forest may take longer than recovery of an area from flooding, and the disamenity of blackened trees may be quite large.

In order to test if the slope coefficients, and hence implicit prices of the hedonic price function changed after the Buffalo Creek fire, we conducted a Chow test of coefficient equality. The Chow test indicated that the coefficients of the hedonic price regressions were significantly different before and after the fire. The calculated F for the linear was 7.52, while for the semi-log was 2.91, while the critical F at the 1% level is 2.64.

## Conclusion

It appears that the Buffalo Creek fire and the additional information the fire conveyed on the risks of living in the wildland urban interface shifted the hedonic price function down. Prospective forest homebuyers took this information seriously, and the rate of increase in forest house prices was reduced by sizeable amounts in the years after the fire. If this finding is replicated in other areas, the good news is that the housing market in the wildland urban interface is starting to reflect the increased hazards of living in these high-amenity forests. This reduction in demand, may, over time, reduce the number of houses built in the wildland urban interface from what it would have been in absence of forest fires.

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**Table 1. Hedonic Property Regression Models**

| | Linear Dependent Variable | | Log Dependent Variable | |
|---|---|---|---|---|
| Variable | Coefficient | t-Statistic | Coefficient | t-Statistic |
| Constant | -1614848. | -7.643 | -8.177116 | -3.966 |
| Garage | 18512.38 | 3.539 | 0.222506 | 4.359 |
| FirePlace | 16073.97 | 4.664 | 0.107021 | 3.182 |
| # of Bathrooms | 21600.63 | 8.056 | 0.171615 | 6.559 |
| YearBuilt | 399.1605 | 4.188 | 0.005560 | 5.980 |
| Acreage | 332.2226 | 5.398 | 0.002667 | 4.441 |
| Trend | 24.60352 | 7.813 | 0.000228 | 7.432 |
| Post-Fire | -17095.53 | -2.131 | -0.160263 | -2.048 |
| Sample Size | 504 | | | 504 |
| Adjusted R-squared | 0.515 | | | 0.502 |
| S.E. of regression | 44820 | | | .437 |
| Mean dependent var | 115747 | | | 11.49 |
| F-statistic | 77.45 | | | 73.305 |
| Probability (F-statistic) | 0.0000 | | | 0.0000 |
| Reduction in House Price After Wildfire | -$17,095 | | | -$18,519 |
| Percent Reduction | -15% | | | -16% |

# Specification of Driving Costs in Models of Recreation Demand

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We thank Jeff Englin, Douglass Shaw, and Scott Shonkwiler for helpful comments. This research was supported in part by the Nevada Agricultural Experiment Station, publication # 51042889.

Specification of Driving Costs in Models of Recreation Demand

Abstract:

Existing recreation demand models have paid much attention to the heterogeneous nature of the opportunity cost of time, but generally stipulate constant per mile costs in access price specifications. This study proposes two alternative approaches to introduce user-specific driving costs into recreation demand models. The first approach is based on a refined measurement of driving costs based on engineering considerations. The second strategy estimates perceived per mile cost as a function of vehicle attributes in an empirical framework. We find strong evidence that driving costs are a visitor-specific concept, and that prescribed and perceived costs differ substantially. However, welfare measures generated by these alternative specifications are not statistically different from those produced by the standard model in our application of jet skiing in the Lake Tahoe region.

## I. INTRODUCTION

The cost of travel to a given destination site is an integral component of virtually all recreation demand models. The specification of travel cost has been the subject of much debate over the last few decades. Traditionally, researchers have interpreted travel cost as an aggregate construct of automotive costs linked to travel distance, the opportunity cost of travel time, and, where applicable, outlays associated with entry fees and other non-discretionary expenses unrelated to time or distance. Since non-discretionary fees are either non-existent or lack sufficient variability to allow for a separate examination of their effect on demand in most empirical settings, they are frequently pooled with automotive costs into a common category, often referred to as "out-of-pocket-costs" (e.g. Ward 1984; McKean et al. 1995; Englin and Shonkwiler 1995a).

As illustrated in Englin and Shonkwiler (1995a), the correct identification and measurement of travel cost components is crucial for an unbiased estimation of visitation demand and welfare effects associated with access restrictions or quality changes at a given site. To date, the debate on travel cost specification has almost exclusively focused on the time component of total cost, both in theory (Ward 1984; Shaw 1992; Shaw and Feather 1999) and in empirical work (McConnell and Strand 1981; Bockstael et al. 1987; McKean, et al. 1995). While existing studies differ in their treatment of time costs, virtually all of them impose arbitrarily chosen vehicle costs per mile[8] in their derivation of out-of-pocket outlays. These per mile auto costs are often derived from external information on vehicle operating costs published by government sources or automobile associations, or simply adopted from other studies. In either case, the same constant per mile cost is assigned to all observations in a given sample of recreationists.

In this study we argue that the automotive component of access cost, just like the opportunity cost of travel time, is a visitor-specific concept, and ought to be modeled accordingly. We propose two alternative approaches to introduce user-specific per mile costs into recreation demand models. The first approach is based on a refined measurement of vehicle costs. It combines information on vehicle operating costs provided by public outlets with directly elicited user-specific information on vehicle type, towing load and passenger composition. These refined "prescribed" vehicle costs are then employed in a multi-site recreation demand framework in lieu of a traditional constant per mile fee. The second approach combines the methodology proposed by McConnell and Strand (1981) and Ward (1983) for an empirical estimation of marginal time costs with the notion of "latent travel price" introduced by Englin and Shonkwiler (1995a) to derive empirical estimates of per mile costs for each respondent. The resulting values are a combination of survey data on vehicle attributes and estimated coefficients. Thus, they can be interpreted as "perceived" automotive costs.

The aim of this study is to (i) examine how prescribed and perceived costs compare to one another, and (ii) to investigate the effect of these alternative specifications of vehicle costs on recreation demand and welfare estimates, especially when contrasted to a standard model with constant per mile fees. We apply our theoretical framework to the analysis of the demand for jet skiing, a recreational activity that has to date received limited attention in the empirical literature. We show that vehicle attributes and combined towing and passenger load have a significant impact on empirical per mile costs, which clearly illustrates the inherently heterogeneous nature of this cost component. While prescribed and perceived vehicle costs differ substantially from one another they do not result in statistically significant differences in welfare measures.

The remainder of this manuscript is structured as follows: in the next section we discuss the theoretical and econometric underpinnings of our proposed demand models. The empirical section of this paper then discusses data, estimation results, and welfare measures. Concluding remarks and a summary of key findings are provided in the last section.

## II. MODEL FORMULATION

As discussed in LaFrance and Hanemann (1989), multi-site demand models consistent with utility theory can be specified through an incomplete demand system (IDS). In this framework, individuals maximize utility and distribute income over a group of goods of interest and a numeraire composite commodity. The resulting Marshallian demands and Slutsky substitution matrix display all desired utility-theoretic properties.[9]

We follow Englin et al. (1998), Shonkwiler (1999) and Moeltner (2003) by combining this theoretic framework with a count data model for seasonal trip demand. Specifically, expected demand by person $i$ for site $j$ is stipulated as

$$E(y_{ij}) = \lambda_{ij} = \exp\left(\alpha + \boldsymbol{\beta}_q' \cdot \mathbf{q}_j + \beta_{p,ij} \cdot p_{ij} + \boldsymbol{\beta}_{p,ik}' \cdot \mathbf{p}_{ik} + \beta_{m,i} \cdot m_i\right) \quad , \qquad [1]$$

where $y_{ij}$ and $\lambda_{ij}$ are the actual and expected demand by individual $i$ for site $j$, respectively, $\alpha$ is a common intercept, $\mathbf{q}_j$ is a vector of site features, $p_{ij}$ is the price of site $j$ to individual $i$ measured as travel cost from $i$'s residence to destination $j$, $\mathbf{p}_{ik}$ is a vector of prices to all other sites in the system, $m_i$ denotes individual income, and the $\beta$-terms are coefficients. Specifically, $\beta_{p,ij}$ is the own-price coefficient for site $j$, and $\boldsymbol{\beta}_{p,ik}$ is a vector containing all cross-price coefficients.

Imposing a valid set of cross-equation restrictions consistent with this semi-log functional form as given in LaFrance (1990), and adding the simplifying constraints of user- and destination-invariant preferences for trail features and user-invariant price coefficients, (1) reduces to

$$E(y_{ij}) = \lambda_{ij} = \exp\left(\alpha + \boldsymbol{\beta}'_q \cdot \mathbf{q}_{ij} + \beta_{p,j} \cdot p_{ij} + \beta_m \cdot m_i\right). \quad [10] \qquad [2]$$

We estimate and test three different models for this analysis. All three models estimate demand of users to the entire system of sites and share the same general expression for price:

$$p_{ij} = \gamma_m \cdot d_{ij} + c_{ij} + \gamma_t \cdot \overline{m}_{h,i} \cdot t_{ij} \quad , \qquad [3]$$

where $d_{ij}$ is the round trip distance for person $i$ to site $j$, $c_{ij}$ are non-discretionary out-of-pocket costs unrelated to distance or travel time, $t_{ij}$ is the round trip travel time in hours, $\gamma_m$ is the automotive component of travel cost per mile, and $\gamma_t$ is a fraction of average hourly wage rate $\overline{m}_{h,i}$. Accordingly, the last term in (3) denotes the opportunity cost of round-trip travel time (OCT). We thus follow the theoretical approach proposed by McConnell and Strand (1981), and recent empirical studies (e.g. Englin et al. 1997; Englin et al. 1998; Train 1998; Moeltner 2003) by specifying OCT to be a function of average hourly wage. As in the cited contributions, we set $\gamma_m = 1/3$.[11] This has been found to be a reasonable approximation to the actual (unobserved) fraction of working wage associated with visitors' valuation of time (Englin and Shonkwiler 1995a). As indicated in (3) we further assume marginal OCT to be constant over time, which is consistent with the empirical finding reported in McKean et al. (1995).

This analysis builds on three different specifications for the vehicle component of (3). Model 1 follows the customary procedure of imposing a constant per mile cost for all respondents, based on aggregate estimates published by automobile associations or government outlets. Some examples for such across-the-board auto cost figures from recent recreation demand studies are \$0.27 (Shonkwiler 1999), \$0.3 (Englin and Shonkwiler 1995a,b; Moeltner 2003), and \$0.38 (Feather and Shaw 1999), where all cost terms are expressed in 2002 dollars. We follow the majority of these studies and set $\gamma_m$ = \$0.3 for our first model. Thus, the full cost component in expected demand for Model 1 is given as

$$\beta_{p,j} \cdot p_{ij} = \beta_{p,j} \cdot \left(0.3 \cdot d_{ij} + c_{ij} + \gamma_t \cdot \overline{m}_{h,i} \cdot t_{ij}\right) \qquad [4]$$

Model 2 refines this specification by utilizing additional observed information on vehicle type, towing load, and passenger load for a respondent specific "best practice" derivation of $\gamma_m$, leading to total cost

$$\beta_{p,j} \cdot p_{ij} = \beta_{p,j} \cdot \left(\gamma_{m,i} \cdot d_{ij} + c_{ij} + \gamma_t \cdot \overline{m}_{h,i} \cdot t_{ij}\right). \qquad [5]$$

To be specific, Model 2 does not estimate $\gamma_{m,i}$ empirically. Instead, it simply utilizes additional visitor-specific information to compute a technically more precise *measurement* for driving cost per mile and thus access price. The detailed computation of $\gamma_{m,i}$ is described in the next section. Model 3, in contrast, is specified to generate *empirical estimates* for per mile cost. It is implemented by parameterizing the distance factor as

$$\gamma_{m,i} = \exp\left(\ln 0.3 + \gamma_0 \cdot g(v_i) + \gamma_1 \cdot h(w_i)\right) \qquad [6]$$

where $g(v_i)$ and $h(w_i)$ are pre-defined functions of vehicle type and total load, respectively, and $\gamma_0$ and $\gamma_1$ are *additional model parameters*. Thus, the model allows for the identification of significant predictors of $\gamma_{m,i}$. This approach is conceptually similar to the one taken by McConnell and Strand (1981) to derive an empirical estimate of the fraction of wage rate to capture the opportunity cost of time (OCT), and by Ward (1983) to directly estimate the OCT. It also corresponds to the notion of "latent price" discussed in Englin and Shonkwiler (1995a) as it uses observed travel characteristics as indicators of some underlying, unobserved cost component in a stochastic framework. Models 2 and 3 allow for a comparison of "best practice" auto cost measurement based on scientific data as generated by Model 2 with estimated per mile costs as perceived by respondents. As discussed in Shiftan and Bekhor (2002) for a sample of daily commuters, best-science and perceived auto costs can diverge significantly. We will examine if this discrepancy also holds for our sample of recreationists.

Model 3 combines expression (6) with (3) and (1) for estimation through a full information maximum likelihood (FIML) procedure. It should be noted that both Models 2 and 3 allow for individual specific mile cost terms, while Model 1 constrains $\gamma_m$ to be equal across observations. Furthermore, by jointly setting $\gamma_0$ and $\gamma_1$ to zero in (6), Model 3 collapses to Model 1. This allows for additional comparative specification tests.

The expected demand specification in equation (1) applies to a randomly drawn member of the relevant population of potential visitors. We assume that this latent demand for a given site follows a type II negative binomial distribution (Cameron and Trivedi 1986), i.e

$$f\left(y_{ij} \mid \mathbf{X}_{\mathrm{ij}}\right)=\frac{\Gamma\left(y_{ij}+v\right)}{\Gamma\left(y_{ij}+1\right)\cdot\Gamma(v)}\cdot\left(\frac{v}{\lambda_{ij}+v}\right)^{v}\cdot\left(\frac{\lambda_{ij}}{\lambda_{ij}+v}\right)^{y_{ij}}, \text{with}$$
$$E\left(y_{ij} \mid \mathbf{X}_{\mathrm{ij}}\right)=\lambda_{ij} \qquad \text{and} \qquad V\left(y_{ij} \mid \mathbf{X}_{\mathrm{ij}}\right)=\left(\lambda_{ij}+\frac{1}{v}\cdot\lambda_{ij}^{2}\right), \qquad [7]$$

where $f$, $E$, and $V$, denote conditional probability density function (pdf), expectation and variance of $y_{ij}$, respectively, $\lambda_{ij}$ is parameterized using observed data $\mathbf{X}_{\mathrm{ij}}$ as shown in (1), $\Gamma$ denotes the mathematical gamma function, and $v$ is the index or precision parameter.

All data for this analysis were collected through on-site interviews of visitors. As is well known, the underlying pdf for trip demands must be adjusted for such cases to avoid estimation bias related to truncation and endogenous stratification (e.g. Shaw 1988; Hellerstein 1992). Englin and Shonkwiler (1995b) examine this issue more closely for models with population demand specified to follow a negative binomial distribution. They show that the density function and first and second moments for trip demand by visitors intercepted on-site corresponding to (7) are given by

$$g\left(y_{ij}^{s} \mid \mathbf{X}_{ij}\right)=\frac{y_{ij}\cdot f\left(y_{ij} \mid \mathbf{X}_{ij}\right)}{E\left(y_{ij} \mid \mathbf{X}_{ij}\right)}=\frac{y_{ij}}{\lambda_{ij}}\cdot\frac{\Gamma\left(y_{ij}+v\right)}{\Gamma\left(y_{ij}+1\right)\cdot\Gamma(v)}\cdot\left(\frac{v}{\lambda_{ij}+v}\right)^{v}\cdot\left(\frac{\lambda_{ij}}{\lambda_{ij}+v}\right)^{y_{ij}},$$
$$E\left(y_{ij}^{s} \mid \mathbf{X}_{ij}\right)=\lambda_{ij}+1+\frac{\lambda_{ij}}{v} \quad \text{and} \quad V\left(y_{ij}^{s} \mid \mathbf{X}_{ij}\right)=\lambda_{ij}\cdot\left(1+\frac{1}{v}+\frac{1}{v}\cdot\lambda_{ij}+\frac{1}{v^{2}}\cdot\lambda_{ij}\right), \qquad [8]$$

where superscript "s" signifies trip information collected from an on-site interview.

As described in the next section, on-site respondents were also asked to report the number of trips to all *other* sites in the demand system. We specify demand to these ancillary destinations to follow the distribution of latent demand shown in (7).[12] Accordingly, the likelihood function to estimate this system of demand equations combines pdf expressions (7) and (8), with $\lambda_{ij}$ parameterized as in (1), and definition of mile cost following either (4), (5), or (6) for the three estimated models.

## III. DATA

The data for this analysis stem from an on-site survey of jet skiers implemented during the summer seasons of 2001 and 2002 at six lakes and reservoirs in the Lake Tahoe region of the central Sierra Nevada. The survey was administered in-person by several interview teams on selected weekdays and weekends between the end of May and the end of August during both seasons. Interview days were approximately evenly distributed across lakes, with equal counts of specific days-of-the-week for each destination. Each respondent was asked to provide information on details for the trip of interception, as well as information on the number of trips taken to the six lakes during the current season up to the interview date, and planned trips yet to be taken throughout the remainder of the current season.[13] Additional information important to this analysis includes technical details on the vehicle and jet ski used for the trip, as well as the gender and age of other household members present at the site. Overall, the survey effort generated 333 valid sets of responses, approximately evenly distributed over the two seasons.[14]

For this analysis, we retain all observations that (i) visited the site for one day only, (ii) traveled less than 300 miles one way to reach the interview site (iii) for whom the visit constituted the sole purpose of the trip. The resulting data set comprises 155 completed questionnaires, yielding a panel of 155 x 6 = 930 trip-counts for the recreation system. Table 1 summarizes some basic lake and trip characteristics for this sample. As can be seen from the table, the largest number of current-season trips are observed for Lahontan and Boca reservoirs. Both destinations offer numerous easy access and launching points, generally free of charge. Distances from visitor origin to destination are comparable across lakes, with means in the 40 to 70 mile range. Another noteworthy feature captured in the table is the ban on 2-stroke engines in effect at Tahoe for both survey seasons. The detailed information on jet skis collected in the survey allows for the identification of such models, and thus of visitors that were affected by the ban.

The measurement of travel time and distances, and the computation of vehicle costs are the central components to this analysis and merit further discussion. To reduce the collinearity between distance and duration of travel that usually plagues travel cost derivations (e.g. Englin and Mendelsohn 1991; Englin and Shonkwiler 1995b) and to allow for a maximum degree of heterogeneity in both measures, we proceed as follows: estimated travel time between origin and interview destination was directly elicited from respondents. These time reports were combined with distances derived using Geographical Information System (GIS) methodology to compute the average travel speed to each site.[15] These speed measures, in combination with GIS generated distances, were then used to estimate travel time to lakes other than the site of interception for all respondents. The resulting values for travel distance and time are employed in all models considered in this analysis.

As indicated in the previous section, Models 2 and 3 incorporate detailed information on vehicle type and load. As a starting point, we use information collected on visitor vehicles to group them into five categories as suggested by the American Automobile Association (AAA 2003). Each category is associated with a baseline per mile operating cost based on estimated expenses for gas and oil, tires, and maintenance. Since our cost refinements focus on gasoline use, we separate this component from the other AAA cost categories and rescale its contribution to per mile costs using regional prices during the survey period, i.e.

$$\gamma_c = \left(\gamma_{c,tire} + \gamma_{c,\text{maintenance}} + \gamma_{c,oil}\right) + p_{WC,t} \cdot g_c \qquad [9]$$

where the $\gamma$-terms refer to per mile costs in dollars, subscript $c$ denotes a specific vehicle category ($c = 1...5$), $p_{WC,t}$ is the regional (West Coast) gas price per gallon for summer season $t$ ($t$

= 2001, 2002), and $g_c$ is the baseline gas consumption in gallons per mile for vehicle category $c$ as reported by the U.S. Department of Energy (D.o.E. 2003).

Using anthropometric data on a person's weight based on gender and age provided by the Federal Aviation Administration (FAA 2003) in conjunction with survey information on travel group size and composition, we then derive an approximated passenger load for each observed trip. This load is augmented by the weight of a towed jet ski, if applicable, using an industry average of approximately 594 lbs per craft. Excess load beyond the vehicle weight implicitly captured in the AAA table is then defined as total load (passengers + jet ski) minus weight of driver (approximated as 175 lbs as per FAA).

To translate excess load into added per mile costs in Model 2, we employ published information on the relationship of weight and fuel consumption (Louisiana Energy and Environmental Resource and Information Center 2003) to compute an added gas-per mile penalty for each increment of 100 lbs of excess weight. Using regional gasoline prices for the survey periods, this figure is then converted into added dollar costs per mile and 100 lbs load. Thus, the refined auto cost per mile for a visitor using a vehicle associated with category $c$ and showing an excess load of $w_i$ can be expressed as

$$\gamma_{m,i} = \gamma_c + g_{c,w} \cdot \left( w_i / 100 \right) \cdot p_{WC,t} \qquad [10]$$

where $g_{c,w}$ denotes the amount of gasoline per mile required by a vehicle in category $c$ to move 100 lbs of excess weight. This refined measure of auto costs is used in Model 2 in lieu of the constant cost per mile amount employed in Model 1. Model 3 in turn, is implemented by setting $g(v_i) = v_{SUV}$ and $h(w_i) = w_i/100$ in equation (6), where $v_{SUV}$ is an indicator variable equal to one if a given vehicle is a Sports Utility Vehicle (SUV, category 4) and zero otherwise.[16]

Table 2 summarizes auto cost components and sample characteristics for each vehicle category. The table reflects clearly that our sample is dominated by SUV-type vehicles, which comprise close to 85% of all observed car types. This category also exhibits the largest number of towed jet skis and the largest mean for excess load. Not surprisingly, however, mean passenger load is larger for vans (including minivans) than for SUVs. Baseline per mile costs as defined in (9) and published by AAA (2003) are given in the fourth-to-last column. They generally lie in the \$0.1 to \$0.2 range, and are highest for luxury cars, followed by SUVs and large sedans. The last three columns in the table capture summary statistics for derived total vehicle costs per mile as per equation (10). The resulting values indicate considerable cost heterogeneity across categories as well as within a given vehicle class. As is clear from comparing AAA baseline costs to mean prescribed total costs, the load penalty significantly increases mean cost figures for all categories to values in the \$0.3 to \$0.6 range. It is especially noteworthy that the mean value for the leading category (SUVs) is about twice the magnitude of the standard assumption of \$0.3 / mile.

The set of regressors used in the estimation of Models 1-3 thus comprises the following variables: a constant term ("constant"), lake elevation in 1000 feet ("elevation"), lake surface in units of 10 square miles ("surface"), an indicator variable equal to one if a given lake does not offer jet ski rental and the interviewed visitor does not own a jet ski ("no_rent"), an indicator variable equal to one for the combined outcome of "respondent owns a jet ski banned at Tahoe" and "site = Tahoe" ("ban"), own-site prices for each of three pairs of lakes labeled as "price1", "price2", and "price3" (see below), the SUV indicator described above ("suv"), excess weight ("load"), and household income in log-form ("income").

## IV. ESTIMATION RESULTS

The estimation results for Models 1-3 are given in Table 3. To reiterate, Model 1 defines vehicle costs as per equation (4) with a constant \$0.3 per mile cost assigned to all observations. Model 2 replaces this constant fee with derived individual per mile costs as described above and indicated in equation (5). Model 3, in turn, generates observation-specific empirical estimates for vehicle costs per mile by following cost specification (6) with $g(v_i)$ and $h(w_i)$ expressed as SUV-indicator and excess load, respectively. We will refer to these models as "standard", "prescribed", and "empirical" in the following discussion.

To guard against misspecification of the precision parameter $v$ in (7) and (8), all three specifications are estimated using robust standard errors applying the methods suggested by White (1982) and Gourieroux et al. (1984). Originally, each model was estimated with site-specific own-price coefficients as stipulated in equation (2). A series of preliminary estimation runs and robust Wald tests revealed that price coefficients are not statistically significantly different from each other for some sets of sites. Specifically, we identified three pairs of sites that each share a common price coefficient: Boca and Tahoe (sites 1 and 5), Donner and Lahontan (sites 2 and 3), and Stampede and Topaz (sites 4 and 6). The corresponding price terms are labeled as "price1", "price2", and "price3" in Table 3. All remaining regressors are defined in the previous section. The value of the log-likelihood function at convergence is given in the row following the estimation results.

Generally, all three models fit the underlying data well, with the majority of estimated parameters significant at 5% or higher. The precision parameter for the negative binomial distribution emerges as highly significant, indicating the presence of over-dispersion in the data.

Implicitly, this result supports our choice of an underlying pdf for visits that allows for mean-variance inequality. All price terms have the expected negative sign. They are identical to the third decimal for Models 1 and 3. This numerical equality is a direct result of the similar per mile costs for the two specifications, as shown below. In comparison, price coefficients for Model 2 are clearly lower in absolute terms for two of the three site pairs, an expected result given the higher marginal vehicle costs and correspondingly higher total travel costs specified for this model (see below). Overall, the strongest price effects are found for Donner and Lahontan (price2), followed by the Stampede-Topaz pair (price3). In contrast, the estimated price coefficient for Boca and Tahoe is considerably smaller in magnitude and only marginally significant in Models 2 and 3.

Income effects exhibit the expected positive sign, but are statistically undistinguishable from zero in all models.[17] Elevation has a negative effect on site visits, while a larger surface is considered a positive lake characteristic by our sample of visitors. The first results is likely related to undesirable lower water temperatures at higher elevation, while the second finding is probably associated with more room to maneuver a jet ski and less congestion at larger lakes. As expected, the no-rent indicator has a highly significant large and negative effect on site choice. In contrast, possession of a jet ski model banned at Tahoe ("ban") does not emerge as a significant determinant of trip demand to that destination. There are two possible explanations for this result: i) these visitors bring a different jet ski on trips to Tahoe, or borrow or rent a jet ski on site, or ii) some of these respondents are not aware that their jet ski model is banned at that destination.[18]

Another important result depicted in Table 3 is the emergence of significant mile-cost parameters in Model 3. Clearly, visitors' trip behavior reflects a sensitivity to vehicle weight as indicated by the highly significant, positive coefficient for "load". The significant, negative coefficient for "SUV" is somewhat puzzling: SUV owners behave as if their per mile cost was reduced compared to other vehicle categories. This clearly contradicts the industry-specific information captured in Table 2, where SUV's are associated with the lowest gas mileage and the second highest baseline per mile costs. It is likely that the indicator variable "SUV" absorbs other latent, cost-reducing components of access price. As suggested in Englin and Shonkwiler (1995a), these unobserved factors may include enjoyment of scenery, perceived benefits related to group travel; or simply positive utility derived from operating a comfortable vehicle in a non-urban environment. In addition, it is well possible that SUV drivers perceive the extra load of a jet ski to be less of an inhibition to smooth and speedy travel, compared to visitors with less powerful vehicles. Generally, the significant effect of these indicators for latent mile cost in Model 3 illustrate that an empirical modeling of auto costs can be a promising avenue towards further refinement of travel cost specification.

A series of specification tests was conducted to provide a more rigorous assessment of the appropriateness of the three models given the underlying data. Test results are given in Table 4. The upper part of the table identifies the pair of models to be compared, the type of test, and statistical test results. The bottom half offers a verbal interpretation of test results. Models 1 and 2 are not nested, but share some common covariates. This calls for Vuong's (1989) 2-step test for overlapping specifications. As can be seen from the table, the test rejects model equivalence, but fails to assign superiority. A similar test for Models 2 and 3 fails to reject equivalence in the first place. Thus, based on this test result, Models 2 and 3 fit the underlying data equally well. In contrast, Model 3 is clearly superior to Model 1, as indicated by the likelihood ratio test in the sixth row of Table 4. This result complements the high levels of significance for the coefficients for "load" and "SUV" discussed above.

## V. TRIP PREDICTIONS

We also compare the three models based on their ability to reproduce sample results for trips to each site. This is accomplished by first computing predicted trips for each destination and person, using (1) for trips other than the site of interception and the first moment of (8) for trips corresponding to on-site interviews, and then averaging over sites. The resulting average-trip-per-person counts for all visitors ("all"), intercept observations ("on-site"), and observations associated with sites other than the interview location ("off-site") are given in Table 5. By simple visual inspection all three models generate very similar trip predictions and reproduce sample averages fairly accurately for the entire set of observations ("all"-category). Specifically, both sample statistics and model predictions indicate that the average surveyed jet skier takes between one and three trips to the each of the six destinations per season. The table clearly illustrates the pronounced difference in expected trip counts for on- and off-site observations. Generally, trip counts corresponding to off-site locations are less than one trip per season for all destinations, while the set of on-site observations produces much higher counts of expected trips, ranging from 3.5 (Donner Lake) to over 16 visits per season (Lake Tahoe). Except for trips associated with Tahoe, all on-site counts predicted by the three models are close in magnitude to sample statistics. Off-site counts are slightly under-predicted for sites 2 to 4, and over-predicted for sites 5 and 6, but are still in the same general numerical range as sample counts. Overall, this comparison of expected to actual counts illustrates again the importance of adjusting our underlying count data pdf for on-site observations. Clearly, visitation probabilities and expected trip counts are substantially higher for person-site combinations associated with an on-site interview. In our case, expected counts for on-site observations exceed their off-site counterparts by a factor of 10 to 20 for all destinations.

## VI. COMPARISON OF DRIVING COSTS

A central focus of this study lies in the examination of differences in per mile costs associated with the three models. Table 6 provides an overview of driving cost components and their statistical properties for the three specifications. Details on the derivation of these values are given in the Appendix. Since out-of-pocket-fees and time cost are specified identically for all three models (see equal summary statistics in "fees+time" row of Table 3), differences in access costs are essentially driven by different marginal costs per mile. As indicated in the last row of the upper half of the table, the sample mean of empirical per mile costs generated by Model 3 is located relatively close to the constant \$0.3/mile fee employed in Model 1 for the representative visitor. Furthermore, marginal driving costs produced by Model 3 generally do not deviate far from this benchmark over respondents as indicated by the relatively small sample standard deviation of approximately \$0.12/mile. This translates directly into similar auto costs and total costs for the two models. In contrast, the vehicle cost refinements implemented in Model 2 result in a significantly higher mean per mile cost of \$0.61 for the representative visitor.

The moderate sample standard deviation of $0.21/mile suggests only limited overlap with per mile costs in the other two specifications.

To allow for a more rigorous comparison between the sample means for per mile costs for Models 1 and 2, and the estimated mean of per mile costs for Model 3, we derive the asymptotic standard deviation of the sample mean and compute 95% confidence intervals for the underlying population mean for Models 2 and 3. The standard deviation and the lower and upper bounds of this interval are denoted as std, LB and UB, respectively, in the lower half of Table 6. As can be seen from the table, confidence intervals for both Models 2 and 3 are fairly tight, and clearly do not overlap. In contrast, the imposed per mile cost of Model 1 lies well below the interval for Model 2, and well within the interval for Model 3. This provides statistical evidence that (i) arbitrarily imposed and empirically estimated marginal vehicle costs converge, and (ii) prescribed costs are significantly higher than both imposed and perceived costs in our application. These insights constitute a key finding in this study.

## VII. WELFARE ANALYSIS

To complete the comparison of estimation results generated by our three models, we compute per-trip and seasonal welfare measures for each specification and site. From a policy perspective, these welfare measures can be interpreted as the economic impact on jet skiers of potential further restrictions on jet ski use in the central Sierra Nevada. Specifically, ecological considerations may dictate bans on some or all jet skis at lakes other than Tahoe in the near future. Furthermore, several of the sites included in this analysis are water reservoirs of regional importance. As it has occurred in the past, water needs during drought periods may require drawing down water reserves to levels that make it impossible to launch jet skis and other vessels, leading de facto to a temporary site closure for all motorized use. An increased awareness of the associated welfare changes to recreational users may aid water planners in deriving socially efficient water management strategies.

Given the marginal role of income effects in all three models, exact welfare measures (compensating variation and equivalent variation) will be numerically close to consumer surplus. We thus limit a discussion of welfare results to consumer surplus (CS). In addition, we follow Englin and Shonkwiler (1995b) and focus on latent population surplus rather than surplus to on-site visitors. In other words, we employ the standard expressions for seasonal and per-trip CS associated with the negative binomial pdf (equation 7) to generate welfare measures for the underlying population of users.[19] Thus, expected seasonal and per trip CS for a specific individual and site are given by

$$\begin{aligned} E(CS_{ij}) &= \int_{p_{ij}}^{\infty} \hat{\lambda}_{ij}\, dp = -\frac{1}{\beta_{p,j}} \cdot \hat{\lambda}_{ij} \\ E(CS_{ij} / trip) &= \frac{E(CS_{ij})}{\hat{\lambda}_{ij}} = -\frac{1}{\beta_{p,j}} \end{aligned} \qquad [11]$$

where $\hat{\lambda}_{ij}$, the predicted number of trips made by person $i$ to site $j$, is derived using equation (1) (e.g. Hellerstein and Mendelsohn 1993; Englin and Shonkwiler 1995b).

Resulting estimates are shown in Table 7. As in Shonkwiler and Hanley (2003), we use the bootstrap method to generate statistics of central tendency and 95% confidence intervals for all welfare measures[20]. Since division by the exceedingly small price coefficient for sites 1 and 5 ("price1") in the expressions given in (11) produces excessive outliers for welfare means, we report instead median values for all welfare statistics. Furthermore, since the expected number of trips per season is generally less than one when all visitors are treated as off-site cases (see Table 5), seasonal welfare is slightly smaller than per-trip CS for all destinations. Generally, median welfare measures flowing from Model 2 are higher than those generated by Models 1 and 3 for all sites, while the latter two models produce again very similar results for most destinations. This is expected given the enhanced role of estimated price coefficients in the derivation of welfare expressions, as illustrated in (11), and the relatively smaller absolute magnitude of price effects for Model 2. However, there is substantial overlap in confidence intervals for all sites and models. In other words, the different specifications of per mile cost do not result in *statistically significant* differences in welfare measures for the three models. This constitutes the second key finding flowing from this analysis.

As expected, Lake Tahoe generates the largest welfare statistics, with both per trip and seasonal CS in the \$200-\$250 range.[21] Popular Boca Reservoir is associated with the second highest seasonal welfare effects (\$150-170). The remaining sites generate clearly smaller welfare measures, ranging from \$40 to \$60 for per-trip CS and from \$10 to \$40 for seasonal welfare. Overall, these results appear of reasonable magnitude when compared to welfare measures for other outdoor activities reported in the literature. Specifically, CS per trip appears to be 50-100% higher than the per-trip welfare generally found for hiking and backpacking (e.g. Englin and Shonkwiler 1995b; Moeltner 2003), and of comparable magnitude to per-trip CS associated with related forms of water recreation (e.g. Englin and Shonkwiler 1995a; Loomis and Walsh 1997).

# VIII. CONCLUSION

This study explores alternative specifications for the automotive component of total travel costs in the context of recreation demand analysis. Specifically, we propose two approaches that allow for heterogeneity in this cost factor, which has traditionally been modeled as an arbitrarily imposed constant fee for all visitors. The first approach is based on a more accurate measurement of driving costs, using respondent-specific information on vehicle type and load. This refined cost segment is then added to remaining portions of total access cost to produce a "best practice", or "prescribed" travel cost variable. The second strategy employs similar information on vehicle attributes, but allows each of them to be associated with its own marginal impact on per mile cost in an empirical framework.

The statistical significance of these impacts in our estimation results implicitly confirms the main hypothesis underlying this analysis: just like the opportunity cost of time, driving costs are a visitor-specific concept. The second key finding in this study is that prescribed and empirically derived per mile costs differ substantially from one another, with the former exceeding the latter by a factor of two for the average visitor. Thus, visitors behave as if their individual per mile costs were much lower than would be dictated by engineering considerations. This implies that travelers are either unaware of true driving costs, or that there are other unobserved factors associated with driving that have a cost-decreasing effect. The former argument would imply that people make sub-optimal trip decisions, and that there is a potential for economic efficiency gains through provision of better information on driving costs to households. The latter argument would call for a more complete specification of "latent" per mile cost in our Model 3, perhaps including trip attributes related to scenery, vehicle amenities and details on passenger composition. It is likely that both information and latent price effects are present in our application, and potentially in any recreation demand model that includes driving costs as a component of access price. Disentangling and measuring these effects could pose a challenging, but potentially rewarding task for future research.

We further find that the popular \$0.3 / mile value for vehicle costs is a reasonable approximation for the average visitor in our application, based on a comparison of this constant fee with the statistical properties of per mile costs generated by the empirical model. However, it should be noted that the "average" driver in our study is associated with a combined towing and passenger load of over 700 lbs. This is unlikely to be the case for individuals engaging in other popular recreational activities, such as hiking or angling. For those applications, a fee of \$0.3 / mile may actually *overstate* the underlying perceived vehicle costs that drive visitation behavior.

On the other hand, this standard fee may well become a good approximation to "best practice" engineering costs for less load-intensive activities. This can be deduced from Table 2, which indicates prescribed costs in the \$0.3 range for vehicle categories with no or few towed jet skies. However, as long as people do not behave as dictated by prescribed costs, the benefits of using this refined cost measure in recreation demand analysis are at best questionable. From an econometric perspective, the use of engineering costs will lead to biased coefficient and welfare estimates if they do not correspond to the underlying per mile costs that drive visitation behavior. Thus, ironically, an arbitrarily imposed constant per mile fee may actually be less damaging to estimation results than a best practice engineering measure that grossly deviates from costs as perceived by the representative traveler.

While a series of rigorous specification tests clearly rejects Model 1 with flat per mile fee in favor of Model 3 with empirical driving costs, we do not find significant differences in model accuracy with respect to trip predictions by our three specifications. More importantly, a comparison of bootstrap-simulated welfare measures and associated confidence intervals does not reveal statistically significant differences in these estimates across models[22]. Perhaps more than any other finding in this study, this result casts serious doubt on the usefulness of collecting costly "technical" information on vehicle specifications and load from each respondent in a recreation demand study.

Instead, considering the combined findings of this study, a further development of the empirical modeling of vehicle costs appears to be a more promising avenue for future research on travel cost refinements than a more accurate measuring of auto costs based on technological guidelines. Until travelers start to exclusively follow engineering paradigms, individually assessed per mile costs will remain unobserved, and can at best be approximated by available indicators associated with individual visitors' travel experience.

# IX. APPENDIX: DERIVATION OF SAMPLE STATISTICS AND ASYMPTOTIC PROPERTIES OF PER MILE COSTS FOR MODELS 2 AND 3

The sample statistics for per mile costs in Model 3 are computed as follows: First, the model generates an estimate of $\gamma_{m,i}$ for each observation following equation (6). Since the estimated parameters $\hat{\gamma}_o$ and $\hat{\gamma}_1$ are asymptotically normally distributed, $\hat{\gamma}_{m,i}$ follows a lognormal distribution with expectation $E_{\gamma_0,\gamma_1}[\hat{\gamma}_{m,i}] = \exp(\hat{A}_i + 0.5 * \sigma^2{}_{A,i})$ and variance $V_{\gamma_0,\gamma_1}[\hat{\gamma}_{m,i}] = \exp(2\hat{A}_i + 2\sigma^2{}_{A,i}) - \exp(2\hat{A}_i - \sigma^2{}_{A,i})$, where $\hat{A}_i$ is the expression in parenthesis in equation (6) using $\hat{\gamma}_o$ and $\hat{\gamma}_1$, and $\sigma^2{}_{A,i}$ is the estimated variance of $\hat{A}_i$. The sample mean of individual costs is thus given by $\bar{\hat{\gamma}} = \sum_{i=1}^{N} E[\hat{\gamma}_{m,i}] \big/ N$, and the sample standard deviation by $s = \sqrt{\sum_{i=1}^{N} \left(E[\hat{\gamma}_{m,i}] - \bar{\hat{\gamma}}\right)^2 \big/ (N-1)}$, where N is the sample size. The asymptotic standard deviation of the sample mean for Model 3 can be derived as $s(\bar{\hat{\gamma}}) = \sqrt{\frac{1}{N^2}\left(\sum_{i=1}^{N} V(\hat{\gamma}_{m,i})\right)}$. In contrast, Model 2 invokes the Central Limit Theorem and approximates this statistic as $s(\bar{\gamma}) = s/\sqrt{N}$, where s is the sample standard deviation.

## X. NOTES

[1] That is, for both oppose and support contexts $S_{ij} = 1$ corresponds to a higher level of utility or preference $dU_{2i}$ within each category. One may also envision this intuitively as splitting the data (observations) into support (S) and oppose (O) responses, creating two independent datasets of binary strength of preference responses. The support (S) data include all "support" and "strongly support" responses (LS responses 4, 5); the oppose (O) data include all "strongly oppose" and "oppose" responses (LS responses 1,2). The resulting support and oppose datasets are then vertically "stacked", or pooled into a single binary dataset, such that the directional effect of the utility difference on strength of preference is preserved. The result is a pooled dataset incorporating both the oppose and support data vertically stacked. For examples of similar data transformations, see Mazzotta and Opaluch (1995) and Johnston and Swallow (1999).

[2] Johnston and Swallow (1999) present hypothesis tests that establish the presence of preference asymmetries in two-stage stated preference questions, yet fail to provide a consistent approach that would allow one to model respondents' choices in the presence of such asymmetries.

[3] Perhaps the most straightforward approach to this issue would be to apply generalized ordered logit, an approach which relaxes the proportional odds assumption implicit in traditional ordered logit models (cf. US EPA 2002). We eschew this model in favor of an alternative specification (bivariate probit) which formalizes the hypothesized two-stage decisions implicit in a preference asymmetry model. We thank Scott Shonkwiler for suggesting this alternative approach.

[4] As a practical matter, neutral responses make up a very small proportion of the Likert scale data in question. However, some data is nonetheless discarded in estimating the bivariate probit model.

[5] The $\chi^2$ statistic is calculated as $-2[LR_R - LR_U]$, where $LR_R$ is the log likelihood function of the restricted model in which $\gamma_{oi} = \gamma_{si}$, and $LR_U$ is the log likelihood function of the unrestricted model (7).

[6] Recall, the ordered probit model estimates only one parameter estimate per attribute, which applies over the entire continuum of Likert scale responses.

[7] Full results for all models are available from the authors upon request.

[8] We will refer to this cost component alternatively as "per mile cost", "vehicle cost" "auto cost" or "driving costs" in the remainder of this text.

[9] As discussed in Hanemann and Morey (1992) the key condition for implementation of this structure is that prices outside the set of goods of interest are either explicitly included in the analysis, or, if unobserved, assumed not to vary across the sample that generated data on system demand. This notion also extends to unobserved site characteristics in the context of recreation demand.

[10] While these restrictions explicitly rule out cross-price effects in the uncompensated site-specific demand equations, they still allow for substitution between sites through compensated demands. Specifically, as shown in Englin et al. (1998) and Shonkwiler (1999) the Hicksian cross-price effects are non-zero as long as $\beta_m$ is positive.

[11] As discussed in Ward (1983) Bockstael et al. (1987), and McKean et al. (1995) this specification rests on the assumption of a labor market equilibrium for all visitors, i.e. their ability to easily substitute time for money income. As shown in these studies, an alternative treatment of OCT for people with labor market constraints would be to specify time costs as a function of time alone. However, this requires information on respondents' discretionary time, which was not available for this application.

[12] This implies that we model trip demands to different sites as independent for a given individual, a standard assumption in count data modeling (e.g. Englin et al. 1998; Lutz et al. 2000; Moeltner 2003). While this assumption may imply some efficiency losses in model estimation, it keeps the econometric model tractable, and allows us to retain the focus of this study on the specification of auto costs.

[13] Restrictions on survey length preempted collecting trip details for visits other than the one intercepted on-site. Our analysis thus rests on the implicit underlying assumption that a) respondents correctly estimate the number of future trips during the remainder of the season for

each site, and b) relevant trip details (such as vehicle type and group composition) remain largely unchanged over all trips for a given respondent and season.

[14] The survey instrument is available from the authors upon request.

[15] Time reports that seemed implausible when paired with GIS distances were eliminated from the set of observations used to compute average speed to each lake. The derived speed averages fall into the 50 to 60 mph range for all sites and thus appear to be of reasonable magnitude.

[16] The remaining cost component in equ. (3), $c_{ij}$, captures entrance, launching, and rental fees. These fees generally vary by site (and for some destinations by access point within a given site), and visitor. A problem with assigning rental fees to a given person and trip is the variability in rental time observed for our sample. For this analysis, we simply assume that a given person always chooses the same rental duration for a one-day jet ski outing as the one observed at the interview site.

[17] Presumably, household income is a much more important determinant for acquiring a jet ski and related gear than for undertaking jet ski outings, conditional on possessing the necessary equipment.

[18] Naturally, explanation (i) would contradict our assumption of invariant respondent characteristics (including equipment) over all non-observed trips. However, the resulting

measurement errors should be minor if a different jet ski is used for Tahoe, since the jet ski component of the load variable in Models 2 and 3 is based on an industry average for weight.

[19] This equal treatment of on- and off-site observations rests on the assumption that visitors intercepted on-site are not systematically different from the remaining population of interest in key attributes relevant to our modeling, such as group size, vehicle type, and passenger load. Such difference is not indicated by basic sample statistics.

[20] Specifically, we draw R=1000 sub-samples from our original sample with replacement. All sub-samples are of the same size as the original data set. For each sample, we re-estimate all three models and, for each model, compute the median consumer surplus over all observations. We then extract the mean and the 2.5$^{th}$ and 97.5$^{th}$ percentile of the set of R medians and report the associated values in Table 7. The theoretical underpinnings of the bootstrap method are discussed in Davison and Hinkley (1997).

[21] Naturally, the limited significance of the underlying price coefficient for this site and the commensurate wide spread of the resulting confidence interval for welfare measures call for caution in the interpretation of this result. The same caveat holds for site 1 (Boca Reservoir).

[22] We also estimate our set of models for different specifications of the fraction of hourly wage that enters the computation of opportunity cost of time. While price coefficients adjust accordingly, the general findings summarized in this study hold: prescribed driving costs are distinctly different from perceived costs, and welfare measures flowing from the three models

are of comparable magnitude. These additional estimation results are available from the authors upon request.

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Table 1: Sample Characteristics by Site

| Lake | total trips (current season) | One-way distance (miles) minimum | mean | maximum | elevation (feet) | surface (squ. miles) | Jet ski rental | Jet ski restrictions[b] |
|---|---|---|---|---|---|---|---|---|
| Boca | 358 | 7.2 | 44.5 | 218.8 | 5700 | 1.5 | no | no |
| Donner | 160 | 2.0 | 49.4 | 207.2 | 5969 | 1.5 | yes | no |
| Lahontan | 399 | 4.0 | 59.8 | 261.2 | 4150 | 23.2 | no | no |
| Stampede | 160 | 8.7 | 52.2 | 220.3 | 5949 | 5.4 | no | no |
| Tahoe | 198 | 4.4 | 53.2 | 250.0 | 6230 | 190.8 | yes | yes |
| Topaz | 296 | 1.7 | 66.8 | 227.2 | 5012 | 4.4 | yes[a] | no |
| Total | 1571 | | | | | | | |

[a] Rental at Topaz started in the 2002 season

[b] A ban on 2-stroke engines was implemented at Tahoe in 1999 for environmental reasons

Table 2: Sample Characteristics and Auto Costs by Vehicle Category

| Category | label | number of vehicles | number of jet skis | passenger load (mean, lbs)[b] | tot. excess load (mean, lbs)[b] | miles per gal. | $\gamma_c$[c] (mean,$) | $\gamma_i$[d] min,$ | mean,$ | max,$ |
|---|---|---|---|---|---|---|---|---|---|---|
| 1 | medium car | 3 | 0 | 116.7 | 116.7 | 33.0 | 0.115 | 0.115 | 0.167 | 0.211 |
| 2 | large car | 1 | 0 | 405.0 | 405.0 | 27.0 | 0.130 | 0.350 | 0.350 | 0.350 |
| 3 | luxury car | 3 | 1 | 58.3 | 256.3 | 25.0 | 0.148 | 0.148 | 0.299 | 0.553 |
| 4 | SUV | 131 | 108 | 264.9 | 754.4 | 23.0 | 0.139 | 0.139 | 0.644 | 1.104 |
| 5 | van | 13 | 6 | 371.2 | 645.2 | 25.0 | 0.127 | 0.267 | 0.513 | 0.815 |
| 6[a] | missing | 4 | 2 | 202.5 | 499.4 | 26.6 | 0.132 | 0.132 | 0.417 | 0.757 |
| total | | 155 | 117 | | | | | | | |

[a] Vehicle type was not reported by 4 respondents. For these observations, category averages were employed.

[b] excluding driver (approx. 175 lbs)

[c] see equation (9) in text

[d] see equation (10) in text

Table 3: Estimation Results

| Variable | Model 1 (standard flat per mile cost) Coeff. | s.e. | | Model 2 (prescribed indiv. per mile cost) Coeff. | s.e. | | Model 3 (empirical per mile cost) Coeff. | s.e. | |
|---|---|---|---|---|---|---|---|---|---|
| constant | 2.279 | (0.711) | *** | 3.275 | (0.718) | *** | 2.516 | (1.049) | ** |
| elevation (1000ft) | -0.056 | (0.013) | *** | -0.062 | (0.014) | *** | -0.059 | (0.014) | *** |
| surface (10 sq.miles) | 0.035 | (0.018) | ** | 0.040 | (0.017) | ** | 0.035 | (0.017) | ** |
| no_rent | -1.619 | (0.581) | *** | -1.785 | (0.624) | *** | -1.679 | (0.622) | *** |
| ban | -0.705 | (0.649) | | -0.689 | (0.655) | | -0.656 | (0.681) | |
| price1 | -0.003 | (0.002) | | -0.003 | (0.002) | * | -0.003 | (0.002) | * |
| price2 | -0.023 | (0.005) | *** | -0.015 | (0.003) | *** | -0.023 | (0.006) | *** |
| price3 | -0.015 | (0.004) | *** | -0.010 | (0.002) | *** | -0.015 | (0.004) | *** |
| suv | - | - | - | - | - | - | -0.666 | (0.307) | ** |
| load | - | - | - | - | - | - | 0.085 | (0.031) | *** |
| log of income | 0.061 | (0.080) | | 0.009 | (0.079) | | 0.063 | (0.083) | |
| v | 0.097 | (0.011) | *** | 0.098 | (0.011) | *** | 0.097 | (0.012) | *** |
| Lhf | 936.277 | | | 938.323 | | | 933.2338 | | |

*=sign. at 10%, **=sign. at 5%, ***=sign. at 1%

All models with White-corrected standard errors

n = 930 based on 155 day-visitors

Table 4: Specification Tests

| Test | First model | Second Model | Relationship | Test | Test stat. chi-squ. | Test stat. t | p-value chi-squ. | p-value t |
|---|---|---|---|---|---|---|---|---|
| 1 | model 1 | model 2 | overlapping | Vuong | 23.47 | 0.422 | 0.013 | 0.673 |
| 2 | model 2 | model 3 | overlapping | Vuong | 12.969 | -1.413 | 0.122 | 0.158 |
| 3 | model 1 | model 3 | nested | LR-test | 6.0864 | N/A | 0.0477 | N/A |

| Test | Result (in words) |
|---|---|
| 1 | reject Ho: models fit data equally well, but "winner" indeterminant |
| 2 | fail to reject Ho: models fit data equally well |
| 3 | reject Ho: coefficients on "SUV" and "load" in model 3 are jointly zero<br>i.e. reject model 1 in favor of model 3 |

Table 5: Average Number of Actual and Predicted Trips per Season

| | site | | | | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3 | 4 | 5 | 6 |
| actual | | | | | | |
| all | 2.31 | 1.03 | 2.57 | 1.03 | 1.28 | 1.91 |
| on-site | 9.80 | 3.49 | 9.49 | 4.21 | 16.63 | 10.72 |
| off-site | 0.51 | 0.32 | 0.71 | 0.45 | 0.44 | 0.22 |
| model 1 | | | | | | |
| all | 2.22 | 1.05 | 2.59 | 0.92 | 1.25 | 1.76 |
| on-site | 9.31 | 3.93 | 10.65 | 4.62 | 9.93 | 8.97 |
| off-site | 0.52 | 0.21 | 0.41 | 0.24 | 0.78 | 0.37 |
| model 2 | | | | | | |
| all | 2.16 | 1.22 | 2.64 | 0.86 | 1.27 | 1.81 |
| on-site | 9.12 | 4.76 | 10.85 | 4.32 | 9.94 | 9.13 |
| off-site | 0.48 | 0.19 | 0.42 | 0.23 | 0.79 | 0.40 |
| model 3 | | | | | | |
| all | 2.26 | 1.07 | 2.61 | 0.89 | 1.26 | 1.78 |
| on-site | 9.47 | 4.03 | 10.68 | 4.49 | 9.93 | 9.11 |
| off-site | 0.53 | 0.20 | 0.43 | 0.23 | 0.79 | 0.37 |

Table 6: Comparison of Travel Costs

| | Model 1 (standard flat per mile cost) | | | Model 2 (prescribed indiv. per mile cost) | | | Model 3 (empirical per mile cost) | | |
|---|---|---|---|---|---|---|---|---|---|
| | Sample statistics | | | Sample statistic | | | Sample statistics | | |
| Cost component ($): | mean | median | std | mean | median | std | mean | median | std |
| total | 72.74 | 55.15 | 66.31 | 104.59 | 89.02 | 72.45 | 75.78 | 58.76 | 65.92 |
| auto | 32.59 | 30.24 | 21.19 | 64.44 | 59.51 | 40.67 | 35.63 | 30.14 | 25.39 |
| fees+time | 40.15 | 25.06 | 50.95 | 40.15 | 25.06 | 50.95 | 40.15 | 25.06 | 50.95 |
| auto per-mile | 0.30 | 0.30 | 0.00 | 0.61 | 0.65 | 0.21 | 0.35 | 0.32 | 0.12 |
| | | | | Asymptotic statistics[a] | | | Asymptotic statistics[a] | | |
| | | | | mean | | | mean | | |
| auto per-mile | | | | | | | | | |
| | | | std | 0.007 | | std | 0.006 | | |
| | | | LB | 0.595 | | LB | 0.335 | | |
| | | | UB | 0.623 | | UB | 0.357 | | |

[a] Based on the probability distribution of the sample mean

Table 7: Comparison of Welfare Estimates

| **seasonal CS / person** | | site | | | | | |
|---|---|---|---|---|---|---|---|
| | | 1 | 2 | 3 | 4 | 5 | 6 |
| model 1 | LB | -1544.326 | 3.408 | 11.007 | 8.963 | -1563.261 | 12.267 |
| | median | 157.998 | 9.284 | 23.855 | 18.857 | 203.531 | 26.892 |
| | UB | 1767.751 | 25.682 | 53.408 | 48.446 | 2150.754 | 81.737 |
| model 2 | LB | -1482.623 | 5.108 | 16.545 | 11.648 | -1423.202 | 18.969 |
| | median | 161.302 | 12.65 | 31.461 | 25.160 | 225.13 | 42.003 |
| | UB | 2034.946 | 28.831 | 57.85 | 50.183 | 1888.164 | 88.911 |
| model 3 | LB | -1468.619 | 3.213 | 10.113 | 7.023 | -1212.514 | 11.018 |
| | median | 172.258 | 9.399 | 23.477 | 17.177 | 225.064 | 28.087 |
| | UB | 1689.404 | 26.143 | 56.665 | 40.573 | 1867.079 | 73.564 |

| **per-trip CS** | | site | | |
|---|---|---|---|---|
| | | 1, 5 | 2,3 | 4,6 |
| model 1 | LB | -2036.524 | 27.641 | 41.089 |
| | median | 230.228 | 42.106 | 65.335 |
| | UB | 2309.874 | 76.406 | 138.754 |
| model 2 | LB | -2007.853 | 44.562 | 64.749 |
| | median | 256.349 | 62.844 | 96.808 |
| | UB | 2559.732 | 94.47 | 161.269 |
| model 3 | LB | -1844.877 | 22.776 | 33.606 |
| | median | 251.573 | 44.349 | 67.795 |
| | UB | 2410.743 | 93.804 | 142.988 |

LB (UB) = 2.5$^{th}$ percentile (97.5$^{th}$ percentile) of bootstrapped distribution

# Does ownership matter? Examining the relationship between property values and privately and publicly owned open spaces, streams and wetlands

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January 9, 2004

## Abstract

Numerous studies have examined the relationship between the sale price of single-family residential properties and amenities such as wetlands, streams, natural areas, golf courses and urban parks. This analysis extends previous research by considering whether open spaces within one-quarter mile of properties in the study area are privately or publicly owned and whether resources, such as streams and wetlands, are on private or publicly owned land. The analysis uses detailed structural, neighborhood, and amenity information for approximately 30,000 single-family residential properties sold in an urbanized area of Multnomah County, Oregon between 1999 and 2001. Results indicate that property values are influenced by the ownership of streams, wetlands, specialty parks, golf courses and cemeteries located within one-quarter mile of properties in the study area. Proximity to publicly owned trails and urban parks is also found to have a statistically significant effect on property values.

[7] This research was supported by a Paid Leave Award from Reed College.

**Does ownership matter? Examining the relationship between property values and privately and publicly owned open spaces, streams and wetlands**

# I. Introduction

The preservation of open space has been a major policy focus for state and local governments, the federal government, and for non-profits such as The Nature Conservancy and the Trust for Public Lands. During 2003, 99 measures were passed in 23 states raising a total of $1.3 billion for land conservation (Trust for Public Land 2003).

The relationship between the sale price of single-family residential properties and open spaces such as golf courses, natural areas, cemeteries, specialty parks, and urban parks, and water resources such as wetlands, rivers, and streams, has been studied extensively. These amenities have been found to influence property values with the effect depending on a property's location relative to the amenity, the type and quality of the amenity and, most recently, whether the amenity can be developed or is permanently preserved.

Some open spaces, for example, golf courses, natural areas, cemeteries, and specialty parks, can be owned by the "public," that is, by federal, state, or local governments. Alternatively, these open spaces can be owned by for-profit or not-for-profit private organizations. Water resources, such as wetlands and streams, may be located on either privately or publicly owned land.

The ownership of open spaces, and the ownership of land on which water resources are located, may generate different impacts on surrounding property values. Public ownership may provide benefits such as access for recreation, flood control, noise and pollution abatement, and desirable views that may not be provided, or provided at the same level, by private owners.

The research presented in this paper extends the existing literature by using the hedonic price technique to examine whether the ownership of natural areas, specialty parks, golf courses and cemeteries by private or public entities has a different effect on the price of single-family residential properties sold between 1999 and 2001 in the part of the city of Portland located in Multnomah County, Oregon. Additionally, this paper compares how the location of water resources, such as wetlands and streams, on public or privately owned property is related to the sale price of single-family residential properties located in proximity to these resources. These questions have important implications for land-use planning, for property tax revenue projections, and for discussions about the "private" or "public" provision of goods with strong public good benefits.

## II. Literature

Numerous studies have used the hedonic price method to estimate the relationship between single-family residential properties and open spaces such as golf courses, urban parks, natural areas, riparian buffers and urban forests (Do and Grudnitski 1995; Lutzenhiser and Netusil 2001; Mooney and Eisgruber 2001; Tyrvainen and Miettinen 2000). The majority of studies find that proximity to an open space increases a property's sale price although some studies have found negative effects for properties located within a certain distance of urban parks (Espey and Owusu-Edusei 2001; Weicher and Zerbst 1973), "public lands" defined as public parks, open access parks, greenways, and private land with conservation easements (Smith et al. 2002), and properties with treed riparian buffers (Mooney and Eisgruber 2001).

Proximity to water resources such as streams, rivers, lakes and wetlands has also been found to influence property values (Doss and Taff 1996; Mahan et al., 2000; Mooney and Eisgruber 2001). Mahan et al. (2000) examine six wetland types in Portland, Oregon as well as proximity to lakes, streams, and rivers. The authors find no difference between wetland types in their preferred model, but they do find that living closer to a wetland, stream or lake will increase a property's sale price. Mooney and Eisgruber (2001) estimate that stream frontage increases property values in their study area, a watershed located in a rural part of western Oregon, by 7%.

Studies have investigated how the condition of marshes and urban streams, as well as the quality of water in proximity to residential properties, influences their value. Earnhart (2001) estimates the benefits from restoring marshes at 16.6% of the median house price while Streiner and Loomis (1996) estimate the benefit of restoring urban streams at 3 to 13% of the mean property price in their study area. Wilson and Carpenter (1999) describe the results of several hedonic studies that estimate a positive and statistically significant relationship between improvements in water quality and property values. Recent research (Leggett and Bockstael 2000; Gibbs et al. 2002; Poor et al. 2001) provides additional evidence that water quality is capitalized into property values.

Accessibility and the potential for open space development have been explored in a series of articles. Cheshire and Sheppard (1995) include information on open spaces with and without public access in their study of two towns in England while Geoghegan (2002), Irwin (2002), Irwin and Bockstael (2001) incorporate "permanent" and "developable" open space variables into their models.

Irwin and Bockstael (2001) examine the relationship between property values and three open space types located within a 400-meter radius of properties: privately owned open space that is developable, privately owned open space that is protected from development, and publicly owned open space. The authors find that an increase in the proportion of land within 400-meters of properties that is publicly or privately held increases the sale price of properties in their instrumental variables model.

Geoghegan (2002) estimates that permanent open space, which includes parks and private land with conservation easements, increases residential land values by a factor of three compared to developable open space, which includes privately owned forests, pastures, and agricultural cropland.

Irwin (2002) considers three types of developable open space: privately owned cropland, pasturelands, and forests, and three types of protected open space: private lands protected from development, non-military open space owned by federal, state or county governments, and military land that is considered to be open space. Irwin concludes that different open space types have different effects on property values with preserved open space having a significantly greater effect on surrounding property values than developable cropland and forests.

## III. Theory

The statistical technique used in this study, the hedonic price method, relates a property's sale price to its structural (S), neighborhood (N) and environmental (Q) characteristics. Assuming that housing choices are the result of utility-maximizing decisions, that prices clear the market, and that the study area represents a single housing market, the price of the $i^{th}$ property location ($P_{hi}$) is represented by equation 1.

$$P_{hi} = P_h(S_i, N_i, Q_i) \tag{1}$$

Researchers have used linear, quadratic, double-log, semi-log, and Box Cox transformations to estimate the hedonic price function (Freeman 2003). Since theory cannot provide guidance on the correct functional form, and since the model likely suffers from missing variables (Cropper et al. 1988), several simple functional forms were estimated with a semi-log model providing the best fit. Natural logs were taken of several explanatory variables (lot square footage, building square footage, age, and median income at the census tract level) since previous research (Geoghegan 2002; Mahan et al. 2000) indicates that the relationship between these variables and the dependent variable is nonlinear.

## IV. Study Area

The study area includes the part of the city of Portland, Oregon located in Multnomah County, an area of approximately 92,150 acres. Approximately 18,400 acres in the study area are classified as publicly owned open spaces and approximately 2,000 acres are classified as privately owned open spaces (Metro Data Resources Center 2003). The study area is highly urbanized with only 4.60% of the study area classified as undeveloped land that is zoned for single family residential use (Hall 2003).

Oregon's statewide land use planning Goal 14 requires the establishment of urban growth boundaries for all cities and metropolitan areas in the state (Oregon Department of Land Conservation and Development 2000). The urban growth boundary for the Portland Metropolitan area is managed by Metro, a regional government that serves 1.3 million residents in 24 cities in the Portland area (Metro 2003). The urban growth boundary encourages compact urban development by preventing sprawl; lots available for residential development are generally small with recent development focused on infill and rebuilding on existing lots.

The city is divided into five quadrants. The Northwest quadrant is divided by the Willamette River, which flows north into the Columbia River. Streets east of the Willamette are labeled "North" while those west of the river are labeled "Northwest" (Figure 1). Johnson Creek, Tryon Creek, Fanno Creek, and the Columbia Slough drain Portland's major watersheds, which are tributaries to the Willamette River.

**Figure 1**

The highly urbanized nature of the study area has "resulted in a loss of habitat for native flora and fauna. The natural habitat that remains is fragmented, has low diversity, and is invaded by exotic species" (BES 2000; 2-4). The loss and degradation of habitat is considered a central factor in the decline of salmonids in the study area. Steelhead and Chinook in the upper Willamette River were listed as threatened under the Endangered Species Act in 1999 (NMFS 2002).

Approximately 56 miles of streams and rivers in the study area were classified as water quality limited in 1998 (Oregon Department of Environmental Quality 1998). The Portland area reach of the Willamette River exceeds standards for toxics, biological criteria, bacteria, and temperature resulting in an overall health rating of marginal to poor (BES 2000). Major sources of pollution in the study area include construction activities, vehicular traffic, leaking sewers, fertilizers and pesticides.

## V. Data

The data set contains structural, neighborhood, location, and amenity information for 30,071 arms-length single-family residential property sales in the study area between January 1, 1999 and December 31, 2001. Sales were screened for recording errors, missing information, and duplicate records. Properties that sold for less than their assessed land value were eliminated under the assumption that these transactions were not at arms-length.

Sale price and structural variables were obtained from the Multnomah County Assessor's Office (2002) with base zoning information provided by the City of Portland, Oregon Bureau of Planning (2002). Prices were adjusted to 2000 dollars using the CPI-All Urban Consumers (Bureau of Labor Statistics 2002). Neighborhood variables were obtained from the Census Bureau (2001) and amenity variables -- open spaces, slope, streams, rivers, and wetlands -- were derived using data produced by the regional government (Metro Data Resources Center 2002a, 2002b).

Amenities on a property include the percentage of the property that has a slope of 25% or greater, the percentage of the property with a stream, and whether the property is sloped and has a stream. Amenities in the area within 1/4 mile of each property include the percentage of publicly and privately owned natural areas, specialty parks, trails, golf courses, and cemeteries and the percentage of wetlands and streams on publicly and privately owned land. Information on the percentage of trails, rivers, and urban parks within 1/4 mile of each property is also included.

Summary statistics for structural and property characteristics are provided in Table 1. The highly urbanized nature of the study area is reflected by the mean lot size of 7,062 square feet. The variable %SLOPE reflects the percentage of a property's lot that has a slope of 25% or greater. This variable was created to capture properties located on hills and buttes in the study area – locations that often have desirable views of the city lights, mountains, or rivers, but may also have higher building and maintenance costs.

*Table 1*

Ten categories of open space are included in the model: private natural area, public natural area, private specialty park, public specialty park, trail, urban park, private golf course, public golf course, private cemetery, and public cemetery. Natural areas are defined as parks where 50% or more of the park is in native or natural vegetation (Waiwaiole 1999). Specialty parks have one dominant use with everything in the park being related to that use, for example, boat ramp facilities (Waiwaiole 1999). Summary statistics for these variables are provided in Table 2. The mean percentage of each open space type is small since the average is calculated using all 30,017 observations. The percentage of urban parks within 1/4 mile of properties has the highest average percentage of 2.16%. Almost 60% of open spaces in the study area are classified as urban parks and approximately 48% of properties in the database are located within 1/4 mile of an urban park.

*Table 2*

## VI. Analysis

Regression results are presented in Table 3.[8] A Breusch-Pagan test for heteroskedasticity using the fitted values of the dependent variable fails to reject the hypothesis of constant variance.

The partial derivative of the hedonic price function with respect to any argument is the marginal implicit price of that characteristic, that is, the additional amount that must be paid for the property to achieve the higher level of the characteristic while holding all other factors constant. The estimated coefficients for the structural variables are statistically significant and

[8] Estimated coefficients for base zoning (low density residential, medium density residential, etc.) and for building characteristics (1 story house with basement, 1 story house with basement and attic, etc.) are not included in Table 3. Full results are available from the author.

positive. Property prices are estimated to increase with larger lots, larger houses, larger garages, additional bathrooms and fireplaces. Older homes are estimated to sell for less – a result that is consistent with the literature (Smith et al., 2002; Mahan et al., 2000). The location variables, quadrant and quadrant interacted with distance from the central business district, are statistically significant. Properties in Northwest, Southwest, Northeast and Southeast Portland are estimated to sell for 68%, 46%, 26% and 12% more, respectively, than properties in North Portland. Property prices are estimated to decrease as the distance to the central business district increases.

*Table 3*

Characteristics of the lot include the percentage of the lot that has a slope of 25% or greater, the percentage of the lot with a stream and whether the lot is sloped and has a stream. The estimated coefficients for these variables are statistically significant with a one percent increase in the percentage of a lot that is sloped increasing a property's sale price by approximately $45 and a one percent increase in the percentage of a lot with a stream decreasing a property's sale price by approximately $647.[9] An increase in the percentage of a lot that is sloped may increase construction costs, but it may also indicate a desirable view. Properties that are both sloped and have a stream are estimated to sell for 10.94% less, or approximately $19,150 less, than properties without these characteristics.

Rivers, as in previous studies (Benson et al. 1998; Kulshresththa and Gillies 1993), are found to have a significantly positive effect on property values. An increase in the percentage of private wetlands or private streams is found to have a significantly negative effect, while an increase in the percentage of public wetlands or public streams has a significantly positive effect. The hypothesis that private and public wetlands have the same impact on property values is

[9] Dollar estimates are based on the mean real sale price of $175,133.

rejected ($F(2, 30{,}017) = 14.23$); the hypothesis that private and public streams have the same effect on property values is also rejected ($F(2, 30{,}017) = 17.98$). The estimated percentage changes and dollar impacts for a one percent change in publicly and privately owned open spaces and water resources located within 1/4 mile of properties in the data set are presented in Table 4.

*Table 4*

Ten categories of open space are included in the model: private natural area, public natural area, private specialty park, public specialty park, trail, urban park, private golf course, public golf course, private cemetery, and public cemetery. The estimated coefficients on all but three open space categories (public natural area, private golf course, and public cemetery) are significant at the 10% level. The estimated coefficients are positive with the exception of trails and private cemetery.

Hypothesis tests were conducted to determine if the estimated coefficients for public and private ownership of natural areas, specialty parks, golf courses and cemeteries are equal. The hypothesis of equal effect was rejected at the 1% level for specialty parks, golf courses and cemeteries, but could not be rejected at conventional significance levels for natural areas. F-statistics and p-values are reported in Table 5.

*Table 5*

Public and private golf courses and specialty parks have a significantly positive impact on surrounding properties, but the impact varies by ownership. The estimated coefficient for public golf courses is over three times larger than private golf courses and the estimated coefficient for private specialty parks is approximately three times larger than public specialty parks. The difference in estimated coefficients likely reflects differences in accessibility for golf courses and, perhaps, qualitative differences for specialty parks.

## VII. Conclusions

The analysis presented in this paper illustrates that the sale price of properties in the study area is affected by both the type and ownership of open spaces and whether wetlands and streams located within 1/4 mile of properties are located on privately or publicly owned land.

The estimated coefficients for publicly owned open space variables are positive although only public golf courses and public specialty parks are statistically significant. The coefficients on the private open space variables yield mixed results -- private cemeteries are significantly negative, private specialty parks and private natural areas are significantly positive, and private golf courses are not statistically significant.

Trails and urban parks located within 1/4 mile of properties in the data set are all publicly owned. The percentage of trails located within 1/4 mile of a property is found to have a significantly negative effect on property values while the percentage of urban park is significantly positive. The trails coefficient may reflect omitted variable bias since many trails in the study area are located near industrial areas.

The coefficients for percentage of privately owned land with wetlands or streams are significantly negative while the coefficients for percentage of publicly owned land with wetlands or streams are significantly positive. These results may reflect differences in accessibility, the current quality of these resources, and beliefs about how the water resources and surrounding property will be managed in the future. Other studies have shown that water quality is capitalized into property values (Leggett and Bockstael 2000; Gibbs et al. 2002; Poor et al. 2001) and that restoring streams in an urban setting can increase property values (Streiner and Loomis 1996). Future research should investigate the extent to which the estimated coefficients reflect qualitative differences, differences in accessibility, and uncertainty over future resource management.

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Figure 1: Study Area

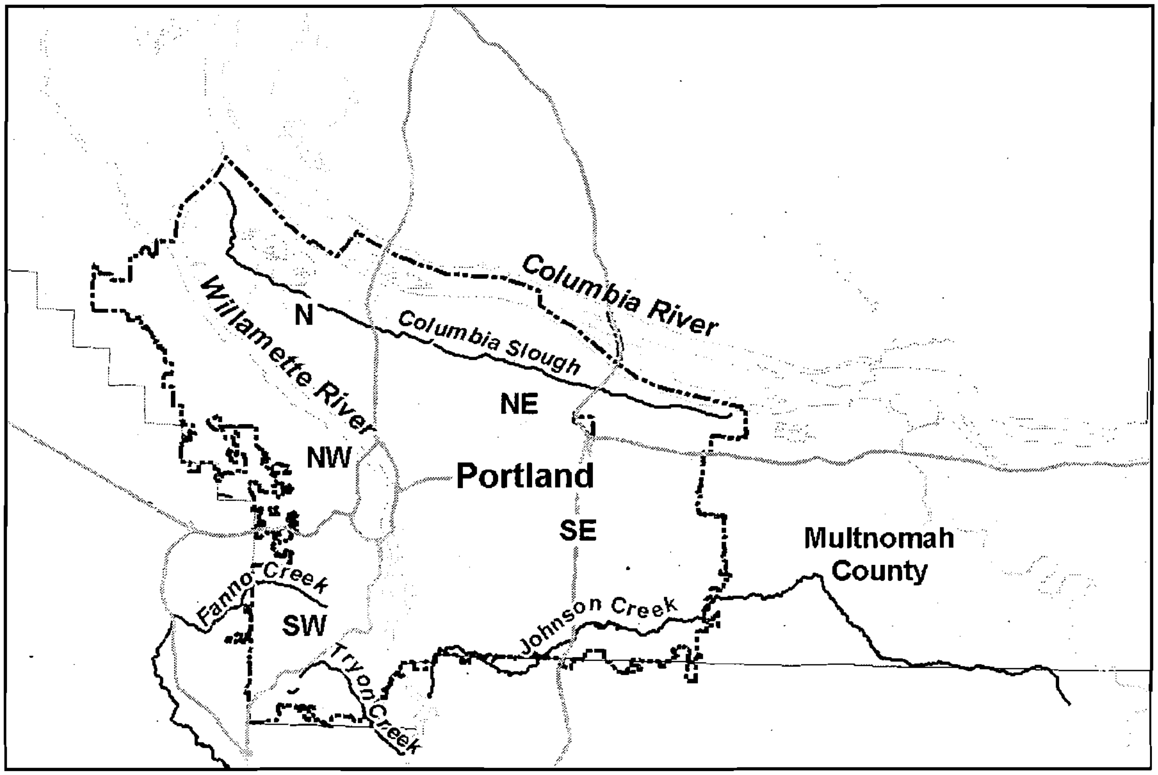

Table 1: Summary statistics: structural, property, and neighborhood variables

| Variable Name | Definition | Units | Mean Value | Standard Deviation | Min | Max |
|---|---|---|---|---|---|---|
| PRICE | Real sale price | $ (2000) | 175,133 | 108,629 | 22,680 | 2,783,203 |
| LOTSQFT | Lot square feet | Square feet | 7,062 | 7,198 | 961 | 324,469 |
| SQFT | Building square feet | Square feet | 1,503 | 691 | 288 | 14,720 |
| GARSQFT | Garage square feet | Square feet | 245 | 205 | 0 | 1800 |
| BATH | Total number of bathrooms | Count | 1.49 | 0.66 | 0.5 | 9 |
| FIREPLACES | Total number of fireplaces | Count | 0.83 | 0.71 | 0 | 8 |
| AGE | Age of building | Count | 59 | 27 | 0 | 155 |
| %SLOPE | Percentage of the property with a slope of 25% or greater | Percent | 2.50 | 13.43 | 0 | 100 |
| %STREAM | Percentage of the property with a stream | Percent | 0.044 | 0.66 | 0 | 27.95 |
| SLP*STRM | Property is sloped and has a stream | Dummy with Yes = 1 | 0.0026 | 0.051 | 0 | 1 |
| INCOME | Median income at the census tract | $ (2000) | 45,974 | 15,446 | 14,091 | 111,064 |
| %WHITE | Percentage white at the census tract | Percent | 77.79 | 13.34 | 29.43 | 95.71 |

Table 2: Summary statistics for open spaces and water resources

| Variable Name | Definition | Mean Value | Standard Deviation | Max |
|---|---|---|---|---|
| %PRWETLAND | Percentage of the area within 1/4 mile of the property with wetlands on private property | 0.044 | 0.44 | 15.81 |
| %PUWETLAND | Percentage of the area within 1/4 mile of the property with wetlands on public property | 0.063 | 0.74 | 26.24 |
| %RIVER | Percentage of the area within 1/4 mile of the property with a river | 0.14 | 1.98 | 57.59 |
| %PRSTREAM | Percentage of the area within 1/4 mile of the property with streams on private property | 0.80 | 0.23 | 1.94 |
| %PUSTREAM | Percentage of the area within 1/4 mile of the property with streams on public property | 0.017 | 0.072 | 1.18 |
| %PRNATURAL | Percentage of the area within 1/4 mile of the property with private natural areas | 0.24 | 2.12 | 29.74 |
| %PUNATURAL | Percentage of the area within 1/4 mile of the property with public natural areas | 0.86 | 4.16 | 68.93 |
| %PRSPECIALTY | Percentage of the area within 1/4 mile of the property with private specialty parks | 0.07 | 0.82 | 36.00 |
| %PUSPECIALTY | Percentage of the area within 1/4 mile of the property with public specialty parks | 0.29 | 1.51 | 47.65 |
| %TRAIL | Percentage of the area within 1/4 mile of the property with trails | 0.11 | 0.64 | 8.65 |
| %URBANPARK | Percentage of the area within 1/4 mile of the property with urban parks | 2.16 | 4.35 | 65.70 |
| %PRGOLF | Percentage of the area within 1/4 mile of the property with private golf courses | 0.085 | 1.51 | 51.62 |
| %PUGOLF | Percentage of the area within 1/4 mile of the property with public golf courses | 0.36 | 2.81 | 73.14 |
| %PRCEMETERY | Percentage of the area within 1/4 mile of the property with private cemeteries | 0.31 | 2.57 | 71.28 |
| %PUCEMETERY | Percentage of the area within 1/4 mile of the property with public cemeteries | 0.09 | 1.08 | 36.55 |

Table 3: Regression results, dependent variable LNRealPrice, N = 30,071, $R^2$ = 0.7781

| Variable | Coefficient | t-statistic | p-value |
|---|---|---|---|
| LN Lot Square Footage | 0.0910061 | 24.07 | .000 |
| LN Building Square Footage | 0.4462804 | 81.41 | .000 |
| Garage Square Footage | 0.0001439 | 21.07 | .000 |
| Total Bathrooms | 0.0735576 | 26.25 | .000 |
| Fireplaces | 0.0420425 | 19.00 | .000 |
| LN Age | -.0061956 | -11.14 | .000 |
| LN Median Income | 0.1407976 | 19.54 | .000 |
| Percentage White | 0.0058409 | 39.50 | .000 |
| Northwest | 0.5157418 | 23.45 | .000 |
| Northeast | 0.2342398 | 16.25 | .000 |
| Southeast | 0.1133889 | 7.58 | .000 |
| Southwest | 0.3796879 | 21.39 | .000 |
| North*CBD | -5.36e-06 | -9.71 | .000 |
| NW*CBD | -1.64e-05 | -18.75 | .000 |
| NE*CBD | -1.16e-05 | -40.54 | .000 |
| SE*CBD | -8.15e-06 | -36.75 | .000 |
| SW*CBD | -1.95e-05 | -37.27 | .000 |
| Percentage of lot sloped | 0.0002561 | 2.30 | .021 |
| Percentage of lot with a stream | -0.003702 | -1.72 | .085 |
| Slope*Stream | -0.1158246 | -4.13 | .000 |
| 1/4 mile percentage of private wetlands | -0.0158478 | -4.10 | .000 |
| 1/4 mile percentage of public wetlands | 0.0098437 | 5.34 | .000 |
| 1/4 mile percentage of river | 0.006261 | 9.64 | .000 |
| 1/4 mile percentage of private streams | -0.0476787 | -6.08 | .000 |
| 1/4 mile percentage of public streams | 0.0421379 | 1.91 | .057 |
| 1/4 mile percentage of private natural area | 0.001741 | 1.84 | .066 |
| 1/4 mile percentage of public natural area | 0.0000781 | 0.19 | .847 |
| 1/4 mile percentage of private specialty park | 0.0057439 | 3.80 | .000 |
| 1/4 mile percentage of public specialty park | 0.0017873 | 2.13 | .033 |
| 1/4 mile percentage of trail | -0.0101161 | -5.03 | .000 |
| 1/4 mile percentage of urban parks | 0.0020057 | 6.86 | .000 |
| 1/4 mile percentage of private golf course | 0.0014717 | 1.34 | .181 |
| 1/4 mile percentage of public golf course | 0.0044913 | 10.02 | .000 |
| 1/4 mile percentage of private cemetery | -0.0023381 | -4.83 | .000 |
| 1/4 mile percentage of public cemetery | 0.0008151 | 0.72 | .474 |

Table 4: Estimated change in real sale price from a one percent change in the water resource or open space

| Amenity type | Estimated change in real sale price (%) | Estimated change in real sale price ($) |
|---|---|---|
| 1/4 mile percentage of private wetlands* | -1.57 | -2,753 |
| 1/4 mile percentage of public wetlands* | 0.99 | 1,732 |
| 1/4 mile percentage of river* | 0.63 | 1,099 |
| 1/4 mile percentage of private streams* | -4.66 | -8,154 |
| 1/4 mile percentage of public streams*** | 4.30 | 7,537 |
| 1/4 mile percentage of private natural area*** | 0.17 | 305 |
| 1/4 mile percentage of public natural area | 0.01 | 14 |
| 1/4 mile percentage of private specialty park* | 0.58 | 1,009 |
| 1/4 mile percentage of public specialty park** | 0.18 | 313 |
| 1/4 mile percentage of trail* | -1.01 | -1,763 |
| 1/4 mile percentage of urban parks* | 0.20 | 352 |
| 1/4 mile percentage of private golf course | 0.15 | 258 |
| 1/4 mile percentage of public golf course* | 0.45 | 788 |
| 1/4 mile percentage of private cemetery* | -0.23 | -409 |
| 1/4 mile percentage of public cemetery | 0.08 | 143 |

*, **, and *** denote significance at the 1%, 5% and 10% levels.

Table 5: Hypothesis tests for equal coefficients

| Hypothesis test for equal coefficients | F-statistic (2, 30,017) | p-value |
|---|---|---|
| 1/4 mile percentage of public and private wetlands | 14.23 | 0.0000 |
| 1/4 mile percentage of public and private streams | 17.98 | 0.0000 |
| 1/4 mile percentage of public and private natural areas | 1.19 | 0.3049 |
| 1/4 mile percentage of public and private specialty parks | 8.57 | 0.0002 |
| 1/4 mile percentage of public and private golf courses | 37.58 | 0.0000 |
| 1/4 mile percentage of public and private cemetery | 10.77 | 0.0000 |

# Hedonic Property Valuation Study:
# Water Chemistry versus Biological Indicators
# for the St. Mary's River Watershed

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## Abstract

The St. Mary's River watershed is located along the western shore of the Chesapeake Bay in the southern part of St. Mary's County, within a sub-watershed of the Potomac River. This research investigates the value that residents within the St. Mary's River watershed place on environmental quality as measured by water quality data from twenty-five monitoring stations located throughout the watershed. No other studies have taken into consideration the effects of ambient water quality within a watershed using the hedonic property valuation framework. A hedonic property value model is used to investigate the marginal implicit values of the following chemical and biological water quality variables: total suspended solids, dissolved inorganic nitrogen, dissolved oxygen saturation, secchi disk depth, Aesthetic Rating, Index of Biological Integrity and Fish Index of Biological Integrity. The results for variables with significant coefficient estimates indicate the marginal implicit prices for the chemical variables, total suspended solids and dissolved inorganic nitrogen, are \$875 and \$18,780, respectively. The marginal implicit prices of the variables with significant coefficient estimates for qualitative measures of stream health and quality, as measured by biological abundance and aesthetic index, are \$1,516 and \$25,225, respectively.

## Introduction

The St. Mary's River watershed is located in St. Mary's County in Southern Maryland. It is considered one of the healthiest tributaries of the Chesapeake Bay. No point sources pollute its waters and its only threat has been run off from local farms and occasional storm runoff. However, over the last few years there has been an expansion of the Patuxent River Naval Air Station; the largest employer in the county. Starting in the late 1990s, the Naval Air Station began to dramatically increase the number of people it employs (MD Dept. of Natural Resources 2002). The rise in employment has lead to the transformation of Southern St. Mary's County. Residential and commercial development along with other associated infrastructure, have been constructed to accommodate the nine percent population increase from 1998 until 2003 (Total Resident Population 2004). This development has lead to an increase in the amount of impermeable surfaces, which in turn, restricts the amount of rainwater percolating into the earth. Restricted absorption results in excess runoff which endangers small streams by overloading them with sediment during storm events.

The Patuxent River Naval Air Station is located at the headwaters of the St. Mary's River watershed where the majority of the streams that feed the St. Mary's River are located. The combination of increased development and numerous small streams, increases the potential for sediment to migrate through the streams and into the river. Biologists monitoring the St. Mary's River watershed warn of the dangers of development and have stated, "In a high discharge year, the water quality will deteriorate, degrading habitat and stressing estuarine organisms of both the watershed and estuary (MD Dept. of Natural Resources 2002)."

County planners are aware of the effects of development and have taken measures to reduce its impact. They have established rural preservation areas, set a minimum for tidal and nontidal wetland buffers and confined development to the major business district along State Route 235 within the town of Lexington Park, where the Patuxent River Naval Air Station is located (MD Dept. of Natural Resources 2002 and Figure 2).

This research study utilizes the hedonic property value model to investigate the value that residents within the St. Mary's River watershed place on environmental quality, using water quality data from twenty-five monitoring stations located throughout the watershed. We develop a hedonic property value model for an entire watershed and associated ambient water quality which is determined almost exclusively from non-point sources of pollution including impervious surfaces and agricultural runoff. The econometric model results are then used to estimate the marginal implicit values of various water quality variables. The following sections include a review of related literature; a review of the hedonic theory; a description of the data; a review of the econometric model; results and conclusions.

## RELATED LITERATURE

Several hedonic property value studies have concluded that water quality does influence waterfront-housing prices (Boyle et al. 1999 and 1998; Epp and Al-Ani 1979; Gibbs et al. 2000; Hsu 2000; Leggett and Bockstael 2000; Michael et al. 1996; Steinnes 1992). The majority of these studies find that the most significant water quality variables are those that can be perceived by the property purchaser and/or those that prohibit the use of water for recreational purposes.

None of these studies consider the overall ambient water quality or ecosystem health within a watershed.

Epp and Al-Ani (1979) studied waterfront residential properties located along small rivers and streams in Pennsylvania and found that pH levels low enough to limit recreational use do affect housing prices. In their study, they tested several water quality variables based on the possibility that they could prohibit recreational use and also be perceived by homeowners. These variables include pH, dissolved oxygen, biochemical oxygen demand, acid from minerals, acid from carbon dioxide, and nitrate and phosphate concentrations. They found that acidity from minerals, acidity from carbon dioxide and in turn pH, significantly influenced housing prices. In addition to testing individual water quality measures, Epp and Al-Ani preformed regressions with a dummy variable that classified streams as clean or polluted based on standards set by the Environmental Protection Agency. Since pH is the most widely used measure of acidity, they included this variable in their final analysis.

A series of studies throughout Maine, New Hampshire and Vermont found water clarity to significantly influence lakefront-housing prices (Boyle et al. 1998; Michael et al. 1996; Gibbs et al. 2002; Hsu 2000). The primary phenomenon affecting water clarity in the studied lakes is eutrophication. Eutrophication is caused by an overload of nutrients that results in an increase in algal growth and eventually leads to reduced water clarity and correspondingly dissolved oxygen. In the Maine lakes study, owners' perceptions of water clarity as well as secchi disk readings were found to be significant throughout the 36 lakes examined in the study (Poor 2001). A similar study, done in New Hampshire found that homeowners are concerned with water quality and that a one meter decrease in visibility lead to a decrease in property value by up to

6% (Gibbs et al. 2002). A third study using Vermont lakes, found that homeowners are willing to pay to prevent a decrease in water clarity and willing to pay to prevent an increase in weed density (Hsu 2000). Boyle et al. (1999) completed both stages of the hedonic model for lakes in Maine that had varying water quality. They found declining marginal values for quantity of water quality associated with lakefront property owners.

Steinnes (1992) studied lakes in Northern Minnesota using continuous water clarity measures as opposed to dummy variables for clean and polluted, as was the case in the Epp and Al-Ani study. Using the value of unimproved residential lots, Steinnes found that water clarity, as measured by secchi disk readings, was positive and significantly related to the sale price.

As previously noted, no studies have investigated ambient water quality or ecosystem health within a watershed. There are however, numerous hedonic property valuation studies regarding ambient air quality within air-sheds (Chattopadhyay 1999, Palmquist 1999, Zable 2000). Each study concluded that there is a significant relationship between the air quality measures, total suspended particulates and sulfur dioxide, and housing prices. The air quality hedonic studies utilize air quality measures from monitoring stations located throughout the airshed. This novel investigation uses a hedonic property valuation model to estimate the marginal implicit prices of water quality data recorded at twenty-five monitoring stations located throughout a watershed in southern Maryland. This approach allows our study to include approximately 1600 residential property sales within a watershed, where only about two percent are actual waterfront properties.

## Hedonic Property Value Model

There are two stages to the hedonic property valuation model. The first stage involves estimating the hedonic price function where the price of a good is regressed on its characteristics to determine the value consumers place on the characteristics that comprise the differentiated good (Feenberg 1980).[10] A house is an example of a differentiated good where its selling price is dependent on its characteristics, where the property sale price represents the market clearing equilibrium price (Rosen 1974). For environmental quality studies, the typical hedonic price function includes housing (S), neighborhood (N) and environmental quality characteristics (E), as denoted as follows:

$$HP = f(S, N, E) \tag{1}$$

The coefficient estimates from a linear regression model are the implicit prices of the characteristics of the differentiated good. In other words, the partial derivative of the hedonic price function with respect to any of the characteristic $i$, gives the implicit price ($\beta$) of that characteristic (Freeman 2001):

$$\frac{\partial P}{\partial Q_i} = \beta_i \tag{2}$$

Where $Q_i$ is the quantity of the characteristic in question. For the hedonic property value model using a semi-log functional form, the marginal implicit price from characteristic '$i$' is calculated as follows:

[10] The second stage of the hedonic model, which will not be carried out in this study, estimates a demand function for characteristics (Taylor 2003).

$$\frac{\partial P}{\partial Q_i} = \beta_i P \qquad (3)$$

Where P is typically equal to the average housing price of the sample used to estimate the coefficient and Q is the quantity of any specific characteristic (Taylor 2003).

The hedonic property value method is appropriate for measuring the value of the benefits of water quality because it allows implicit prices of an unobservable good to be observed through market transactions (Taylor 2003). When a consumer buys a house, he/she is buying a bundle of goods. By looking at the prices of differentiated goods and the amount of each characteristic it possesses the implicit prices of those characteristics can be determined (Freeman 2001). That is, a change in the quantity of a good results in a change in the sales price. The resulting price difference is the implicit price of that characteristic (Taylor 2003). By including measures of ambient water quality or ecological health along with other structural and neighborhood characteristics, the marginal implicit price of the water quality characteristics can be estimated.

## The Study Area and Data

As previously noted, the St. Mary's River watershed is located in the southern part of St. Mary's County along the western shore of the Chesapeake Bay, within a sub-watershed of the Potomac River (Figure 1). It is the largest watershed in St. Mary's County covering approximately 47,000 square acres. The watershed is composed of approximately one hundred miles of freshwater streams and approximately eight linear miles of tidal water. The water quality throughout the watershed varies. Greater amounts of non-point pollution enter from the northern end in proximity of the largest town, Lexington Park. Where as, the southern portion of the watershed consists of primarily rural, farmlands and forests (Figure 1).

### *Housing and Neighborhood Characteristics*

The data set complied for this study includes 1,644 residential sales occurring within the St. Mary's River watershed over four years between June 1, 1999 and May 31, 2003. The property sale prices and characteristics were obtained from SpecPrint Inc., a real estate information collection company that compiles information for the Maryland State Department of Assessments and Taxation. Only arms-length sales of single-family residential properties were included in the data set. Monthly price indices were calculated from June 1999 to May 2003 using county housing price data from Metropolitan Regional Information Systems, Inc (Real Estate Trend Indicator). Using the monthly price indices the property sale prices were converted to January 2003 constant dollars.

In addition to the housing sale price and characteristics, each sale property was geo-coded allowing a GIS dataset to be compiled. Census Block Group median income data was included as a spatial layer. ArcView was used to calculate each property's proximity to Gate One of the Patuxent Naval Air Station and to the landfill located within the watershed. The location of property sales within the watershed, along with the location of the county landfill and Gate One of the Naval Air Station can be seen in Figure 2.

### *Environmental Data*

The water quality data was acquired from twenty-five water-monitoring stations within the watershed. The water quality monitoring activities are part of the St. Mary's River Project conducted by the Biology Department at St. Mary's College of Maryland. The hedonic property model uses the yearly averages of the following chemical and biological water quality variables for each of the monitoring stations: dissolved oxygen saturation, secchi depth, dissolved inorganic nitrogen, pH, total suspended solids, Fish Index of Biological Integrity (FIBI), Index of Biological Integrity (IBI) and an Aesthetic Rating (MD Dept. of Natural Resources 1998). Using the GIS software, each property was assigned to the closest monitoring station, and then was linked by the sale month and year to the corresponding yearly average of the closest monitor's water quality data.

## Econometric Model

Recall the goal of this research is to utilize the hedonic property value model to estimate the implicit values of ambient water quality within a watershed. The specific econometric model used in this research uses the natural logarithm of the real property sale price as the dependent variable and regresses it on a set of structural characteristics (S), neighborhood characteristics (N), and environmental water quality characteristics (E) (Equation 4). The natural logarithm of the real property sale price or the semi-log functional form, has been found to provide a better fit for hedonic data, which was also the case for our dataset (Palmquist 1984).

$$\text{LNREALPR} = \alpha + \beta_{1i}\text{S} + \beta_{2i}\text{N} + \beta_{3i}\text{E} + \varepsilon_i \tag{4}$$

Where $\alpha$, $\beta_1$, $\beta_2$, $\beta_3$, are the coefficients to be estimated. Summary statistics for the explanatory variables are shown in Table 1. The structural characteristics that describe both the physical characteristics of the house include: the age of the house (AGE), number of bathrooms

(BATHS), a garage dummy variable (GARAGE), and the natural logarithm of the building area (LNBLDAR). Explanatory variables describing the land or lot characteristics include the property acreage (ACRES) and a waterfront dummy variable (WTRFRT) which indicates whether the sale property is a waterfront lot. It is expected that the age of the house will be inversely related to sales price, and that the larger the structure, the number of bathrooms, the presence of a garage, number of acres, and if the lot is waterfront, will be positively related to the real price.

A number of neighborhood characteristics were included in the model. They include census tract median income, the distance to Gate One of the Patuxent River Naval Air Station, and distance to the county landfill located within the northern part of the watershed.[11] Distance to Gate One is included to control for distance to the primary employer in the watershed and it is expected that a closer distance would be desirable. In addition, it is expected that proximity of the residential properties to the landfill will have a negative influence on property values. Also, homes located in neighborhoods with higher median income should be positively related to sale price.

The chemical water quality measures included in the model are yearly averages of: the absolute value of the difference from neutral pH (DIFF_PH); dissolved inorganic nitrogen (DIN); dissolved oxygen saturation (DO_SAT); total suspended solids (TSS); and secchi disk readings (SECC_AVG). High concentrations of total suspended solids and dissolved inorganic nitrogen reduce water clarity and impair sub aquatic vegetations' (SAV) ability to grow by blocking light (US Environmental Protection Agency 2003). The growth of SAV is an important

[11] Dummy variables were calculated for high school districts and elementary school districts. The dummies variables were found to be insignificant and omitted from the model.

part of the Chesapeake Bay and its tributaries' food chain and is necessary to protect and guarantee the continuation of the lives of organisms. It is expected that high levels of total suspended solids and dissolved inorganic nitrogen will have a negative impact on sales prices, as this is an indication of compromised water quality associated with non-point source pollution. Since dissolved oxygen is essential to life of aquatic organisms, higher concentrations indicate good water quality and as such are expected to positively influence sales prices.

Aquatic organisms and SAV can only tolerate small changes in pH. Therefore, pH levels higher or lower than the tolerant range are expected to negatively influence property prices. To capture these deviations, the absolute values of the difference from neutral pH measure of 7, were calculated for each sale property. It is expected that the pH coefficient will be negative indicating smaller variation from neutral to be preferred. The secchi depth coefficient is also expected to positive indicating that clearer water is more desirable.

For homes located nearer non-tidal stations, the Index of Biological Integrity (IBI) and the Fish Index of Biological Integrity (FIBI) were included as measures of stream health. These indices measure a streams ability "to support and maintain a biota that is comparable to that found in natural conditions" (Maryland Department of Natural Resources 1998). Microinvertebrates and fish are sensitive to stressors and good indicators of stream health (Maryland Department of Natural Resources 1998). As noted in Appendix A, the IBI for each stream is calculated by comparing the total number and variety of microinvertebrates found at the site with a reference stream, or a stream that is minimally impacted by human activity (Maryland Department of Natural Resources 1998). The calculation of the FIBI is similar to the IBI, except the numbers of fish are counted as opposed to the number of microinvertebrates.

Separate hedonic equations were estimated for the homes located nearer the non-tidal stream monitoring stations, for which the IBI and FIBI data was available. For these equations we expect better ecological health as indicated by higher IBI/FIBI scores to be positively related to the sales prices.

A separate hedonic equation was also estimated for the non-tidal properties using an aesthetic variable (AESTH) which is a subjective measure of stream quality or ecological health, as measured by visual signs of human activity including refuse, remoteness and vegetative state (Maryland Department of Natural Resources 1998; also see Appendix B). It is expected that the Aesthetic Rating will be positively related to sale price.

## Regression Results

As previously noted, in order to minimize multi-collinearity, separate linear regression equations were estimated for each of the water quality variables. The regression results are presented in Tables 2, 3 and 4. The coefficient estimates associated with the variables used to property characteristics (AGE, BATHS, GARAGE, ACRES, WTRFRT, and LNDBLDAR) all had the expected signs. The coefficient estimates for house age, property acreage, and waterfront dummy variables were significantly different from zero in each on the estimated equations. The signs on the coefficient estimates for the distance to Gate One of the Naval Air Station and distance to the county landfill variables were opposite of our initial expectations. One possible explanation for the positive sign on the coefficient estimate associated with the distance to the Naval Air Station lies in the rural nature of this relatively small study area where homebuyers do not necessarily desire to live close to there place of employment, which is located in the most heavily commercial area of the watershed. Rather single family home construction tends to be rather spread out into the rural, less densely populated areas of the watershed, and because traffic volume in the watershed is relatively light, proximity to ones workplace does not seem to outweigh living in more rural areas of the watershed. In addition, many of the residential properties located in Lexington Park are rental apartments which can accommodate the large number of military personnel on temporary assignment, were not included in our data set of single family residential property sales

The landfill in the watershed is located along the same road as one of the largest and most desirable planned communities in this part of the county. Typically, living near a landfill is deemed undesirable because of potential environmental quality issues as well as truck traffic. However, in this case, because the landfill is located in a populated area of the county near a very desirable residential neighborhood, the coefficient estimate on this distance variable has a negative sign indicating as distance to the landfill declines, property values are higher.

Median income was significant in each regression, however its coefficient was opposite our expectations. One possible explanation is that during the time frame selected for this study the housing market within the watershed has been very robust in part because of record low mortgage rates, resulting in consumers buying homes at the higher end of their qualifying price range, leading to the unexpected sign on the coefficient of the income variable.

As previously noted, separate regression equations were estimated for each water quality variable to avoid multicolinearity. As seen in Table 2, the coefficients for total suspended solids and dissolved inorganic nitrogen were of the expected negative sign and were significantly different from zero. Recall that as total suspended solids accumulate within the watershed, water clarity will be reduced. Also recall dissolved inorganic nitrogen directly influences water clarity by stimulating alga growth, which accumulates on the leaves of sub aquatic vegetation and blocks light from reaching them resulting in compromised water quality (US Environmental Protection Agency 2003).

The coefficient estimate for dissolved oxygen saturation has the expected coefficient sign, but is not significantly different from zero (Table 3). The pH coefficient and secchi depth coefficient were not the expected signs, however, they also were not significantly different from zero. The majority of the properties used in this study are not located on properties adjacent to deep waterfront. In fact, most are located near shallow streams and ravines where secchi measures can be limited, thus reducing the number of observations in this regression equation. Therefore unlike the waterfront hedonic studies previously noted, secchi measures are not as good an indicator of the health of ambient watershed water quality or ecological health.

The coefficient estimates for the biological water quality variables along with the aesthetic measure were all of the expected signs, however only the FIBI and the aesthetic rating coefficient estimates are significantly different from zero (Table 4). In other words, healthy stream ecosystems, as measured by the amount and variety of fish species, is a significant and desirable attribute associated with residential properties in the St. Mary's River watershed. The aesthetic variable, which measures visual indicators of stream quality, also has a significant and positive effect on residential housing prices within the watershed.

Marginal implicit prices are calculated for the environmental water quality attributes with significant coefficient estimates (see Table 5). The estimated model indicates that a one unit (mg/L) increase in total suspended solids has a negative impact on average housing prices within the watershed of $875. Correspondingly, a one-unit change (mg/L) in the dissolved inorganic nitrogen, a water quality attribute that manifests itself as a contributor to eutrophication, also has a negative impact on average housing prices in the watershed of $18,780. For the significant biological indices, Aesthetic and FIBI, the empirical results indicate that a one unit increase in each of these indices, results in an increase in average housing prices within the watershed of $1,516 and $25,225, respectively. Recall the Aesthetic Rating scale ranges from 0 to 20 and the FIBI ranges from 1-5 (See Appendices A and B).

## Conclusions

To our knowledge this is the first hedonic property valuation model to estimate marginal implicit prices of ambient water quality and ecosystem health variables for an entire watershed using both primarily non-waterfront properties. Significant ambient water quality variables include both chemical indicators (TSS and DIN) and biological or ecological indicators (ASTHETIC and FIBI).

We have shown that ambient water quality within a watershed can significantly influence residential property values regardless of whether they are waterfront properties. As residential and commercial development continue to intensify within this watershed, it will become important to continue to monitor the ambient water quality, as well as, any potential associated negative impacts on residential property values. Decreasing housing sale prices will be one indicator of the value of the environmental costs, in terms of compromised water quality or ecosystem health, associated with any such development. Given the St. Mary's River Watershed is currently considered one of the least environmentally compromised sub-watersheds of the

Chesapeake Bay watershed, this research helps to quantify the potential future environmental costs associated with development pressures and the associated increases in impervious surfaces, increasing non-point source pollution.

**Table 1 – Variable Descriptions and Expectations**

| Name | Description | Expected Sign | Mean |
|---|---|---|---|
| *LNREALPR* | Log of Real Price (January 2003 constant dollars) | | 12.04267 |
| *AGE* | Age of structure (years) | - | 12.5535 |
| *BATHS* | Number of Bathrooms | + | 1.96612 |
| *GARAGE* | Garage Dummy | + | 0.173358 |
| *ACRES* | Land Area in Acres (calculated to one hundredth of an acre) | + | 1.19235 |
| *WTRFRT* | Waterfront Property Dummy | + | 0.026156 |
| *LNBLDAREA* | Log of the Building Base Area | + | 7.50261 |
| *GATE_DIS* | Distance to Gate One of Patuxent Naval Air Station (meters) | - | 5919.2 |
| *LDFL_DIS* | Distance to Landfill (meters) | + | 7019.44 |
| *MED_INC* | Median income of census block group | + | 52926.2 |
| *DOYEAR* | Dissolved Oxygen Saturation (%) | + | 90.7301 |
| *DIFF_PH* | Absolute value of the difference from neutral pH (standard units) | - | 0.934719 |
| *SECCHI* | Secchi Depth (m) | + | 1.44593 |
| *DINYEAR* | Dissolved Inorganic Nitrogen (mg/L) | - | 0.596739 |
| *TSSYEAR* | Total Suspended Solids (mg/L) | - | 13.0761 |
| *ASTHETIC* | Aesthetic Rating (scale 0-20) | + | 13.4799 |
| *AVG_FIBI* | Fish Index of Biological Integrity (scale 1-5) | + | 4.052871 |
| *AVG_IBI* | Index of Biological Integrity (scale 1-5) | + | 3.32688 |

**Table 2 – Chemical Indicators: Total Suspended Solids and Dissolved Inorganic Nitrogen**

| VARIABLES | TSSYEAR | | DINYEAR | |
|---|---|---|---|---|
| | Coef. (s.e.) | P-value | Coef. (s.e.) | P-value |
| *ONE* | 12.237200* | 0.000 | 12.242000* | 0.000 |
| | (0.283) | | (0.256) | |
| *AGE* | -0.005789* | 0.000 | -0.006363* | 0.000 |
| | (0.001) | | (0.001) | |
| *BATHS* | 0.034576 | 0.269 | 0.024045 | 0.399 |
| | (0.031) | | (0.029) | |
| *GARAGE* | 0.035837 | 0.237 | 0.038278 | 0.158 |
| | (0.030) | | (0.027) | |
| *ACRES* | 0.019824* | 0.000 | 0.020709* | 0.000 |
| | (0.004) | | (0.005) | |
| *WTRFRT* | 0.876694* | 0.000 | 0.875972* | 0.000 |
| | (0.122) | | (0.123) | |
| *LNBLDAR* | 0.021925 | 0.560 | 0.023323 | 0.493 |
| | (0.038) | | (0.034) | |
| *GATE_DIS* | 0.000029* | 0.000 | 0.000036* | 0.000 |
| | (0.000) | | (0.000) | |
| *LDFL_DIS* | -0.000036* | 0.000 | -0.000042* | 0.000 |
| | (0.000) | | (0.000) | |
| *MED_INC* | -0.000003** | 0.014 | -0.000003*** | 0.006 |
| | (0.000) | | (0.000) | |
| *TSSYEAR* | -0.004648*** | 0.060 | | |
| | (0.002) | | | |
| *DINYEAR* | | | -0.101*** | 0.007 |
| | | | (0.037) | |
| SAMPLE SIZE | 1188 | | 1335 | |
| F-STAT | 23.03 (P=.000) | | 26.35 (P=.000) | |
| R-SQUARE | 0.164 | | 0.166 | |

*, **, and *** indicate significance levels of 1%, 5% and 10%, respectively

**Table 3 – Chemical Indicators: Dissolved Oxygen, pH and Secchi Disk**

| VARIABLES | DOYEAR | | DIFF_PH | | SECC_AVG | |
|---|---|---|---|---|---|---|
| | Coef. (s.e.) | P-value | Coef. (s.e.) | P-value | Coef. (s.e.) | P-value |
| *ONE* | 12.051900* | 0.000 | 12.136800* | 0.000 | 12.088700* | 0.000 |
| | (0.291) | | (0.258) | | (0.306) | |
| *AGE* | -0.006448* | 0.000 | -0.006978* | 0.000 | -0.005280* | 0.000 |
| | (0.001) | | (0.001) | | (0.001) | |
| *BATHS* | 0.026837 | 0.351 | 0.029376 | 0.316 | 0.031714 | 0.363 |
| | (0.029) | | (0.029) | | (0.035) | |
| *GARAGE* | 0.035336 | 0.193 | 0.028904 | 0.293 | 0.033751 | 0.306 |
| | (0.027) | | (0.027) | | (0.033) | |
| *ACRES* | 0.021217* | 0.000 | 0.020568* | 0.000 | 0.019666* | 0.000 |
| | (0.005) | | (0.005) | | (0.004) | |
| *WTRFRT* | 0.875189* | 0.000 | 0.951394* | 0.000 | 0.869791* | 0.000 |
| | (0.124) | | (0.110) | | (0.126) | |
| *LNBLDAR* | 0.021234 | 0.535 | 0.015017 | 0.663 | 0.031936 | 0.430 |
| | (0.034) | | (0.034) | | (0.040) | |
| *GATE_DIS* | 0.000034* | 0.000 | 0.000036* | 0.000 | 0.000026* | 0.000 |
| | (0.000) | | (0.000) | | (0.000) | |
| *LDFL_DIS* | -0.000035* | 0.000 | -0.000035* | 0.000 | -0.000033* | 0.000 |
| | (0.000) | | (0.000) | | (0.000) | |
| *MED_INC* | -0.000003** | 0.015 | -0.000002** | 0.035 | -0.000003** | 0.047 |
| | (0.000) | | (0.000) | | (0.000) | |
| *DOYEAR* | 0.001 | .444 | | | | |
| | (0.001) | | | | | |
| *DIFF_PH* | | | 0.029 | 0.315 | | |
| | | | (0.029) | | -0.015 | 0.686 |
| *SECC_AVG* | | | | | (0.037) | |
| SAMPLE SIZE | 1335 | | 1296 | | 1051 | |
| F-STAT | 25.46 (P=.000) | | 27.09 (P=.000) | | 17.6 (P=.000) | |
| R-SQUARE | 0.161 | | 0.174 | | 0.145 | |

*, **, and *** indicate significance levels of 1%, 5% and 10%, respectively

**Table 4 - Estimation Results: Biological Indicators**

| VARIABLES | ASTHETIC Coef. (s.e.) | ASTHETIC P-value | AVG_IBI Coef. (s.e.) | AVG_IBI P-value | AVG_FIBI Coef. (s.e.) | AVG_FIBI P-value |
|---|---|---|---|---|---|---|
| *ONE* | 11.878300* | 0.0000 | 12.159700* | 0.0000 | 11.582200* | 0.0000 |
| | (0.2839) | | (0.2438) | | (0.3599) | |
| *AGE* | -0.008980* | 0.0000 | -0.009567* | 0.0000 | -0.009575* | 0.0000 |
| | (0.0012) | | (0.0010) | | (0.0010) | |
| *BATHS* | 0.018961 | 0.5519 | 0.015330 | 0.5683 | 0.017087 | 0.5295 |
| | (0.0319) | | (0.0269) | | (0.0272) | |
| *GARAGE* | 0.033666 | 0.2381 | 0.027634 | 0.2814 | 0.022551 | 0.3749 |
| | (0.0285) | | (0.0257) | | (0.0254) | |
| *ACRES* | 0.015242* | 0.000035 | 0.017826* | 0.000001 | 0.015641* | 0.000004 |
| | (0.0037) | | (0.0035) | | (0.0034) | |
| *LNBLDAR* | 0.022184 | 0.5591 | 0.012531 | 0.6956 | 0.015397 | 0.6271 |
| | (0.0380) | | (0.0320) | | (0.0317) | |
| *GATE_DIS* | 0.000038* | 0.000020 | 0.000041* | 0.0000 | 0.000046* | 0.0000 |
| | (0.0000) | | (0.0000) | | (0.0000) | |
| *LDFL_DIS* | -0.000021* | 0.0008 | -0.000033* | 0.0000 | -0.000034* | 0.0000 |
| | (0.0000) | | (0.0000) | | (0.0000) | |
| *MED_INC* | -0.000001 | 0.2969 | -0.000002 | 0.1514 | -0.000002*** | 0.0971 |
| | (0.0000) | | (0.0000) | | (0.0000) | |
| *ASTHETIC* | 0.008581** | 0.0250 | | | | |
| | (0.0038) | | | | | |
| *AVG_IBI* | | | 0.001676 | 0.9261 | | |
| | | | (0.0181) | | | |
| *AVG_FIBI* | | | | | 0.136128** | 0.0362 |
| | | | | | (0.0650) | |
| SAMPLE SIZE | 929 | | 1187 | | 1179 | |
| F-STAT | 13.51 ( P = .000) | | 29.32 (P=.000) | | 31.17 (P = .000) | |
| R-SQUARE | 0.1169 | | 0.1831 | | 0.1936 | |

*, **, and *** indicate significance levels of 1%, 5% and 10%, respectively.
Note that only property sales near non-tidal monitoring stations were used in these equations and as such there are no waterfront property sales, and as such this variable is excluded.

**Table 5 - Marginal Implicit Prices**

| Water Quality Attribute | Sample Mean | Coef. Estimate | Average Sample House Price | Marginal Implicit Price |
|---|---|---|---|---|
| TSSYEAR | 13.165857 | -0.004648 | $188,230 | -$875 |
| DINYEAR | 0.62162066 | -0.101 | $185,120 | -$18,750 |
| ASTHETIC | 13.619 | 0.008581 | $176,640 | $1,516 |
| FIBI | 4.053 | 0.136128 | $ 169, 850 | $25,225 |

**Figure 1 - Reference Map**

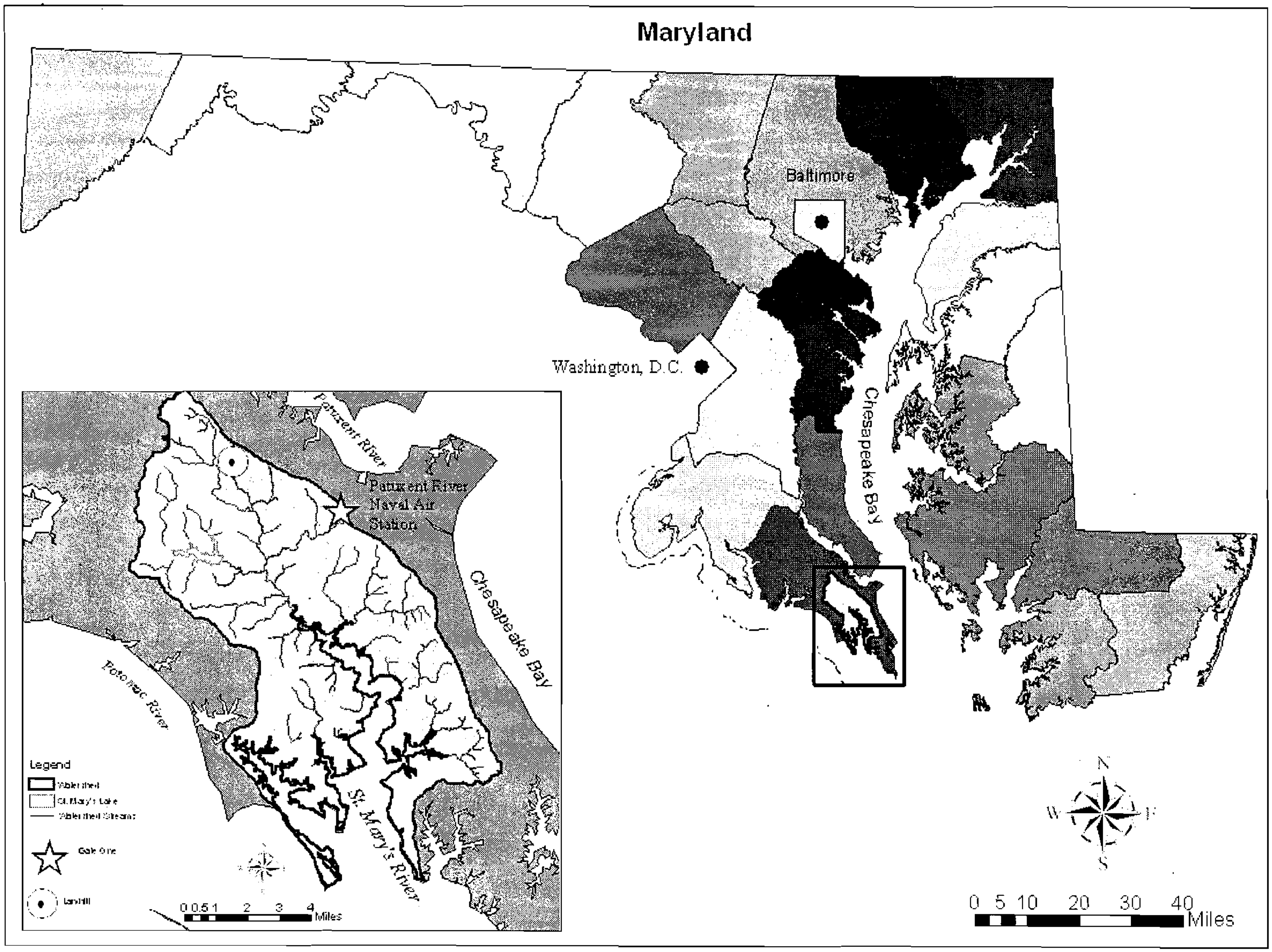
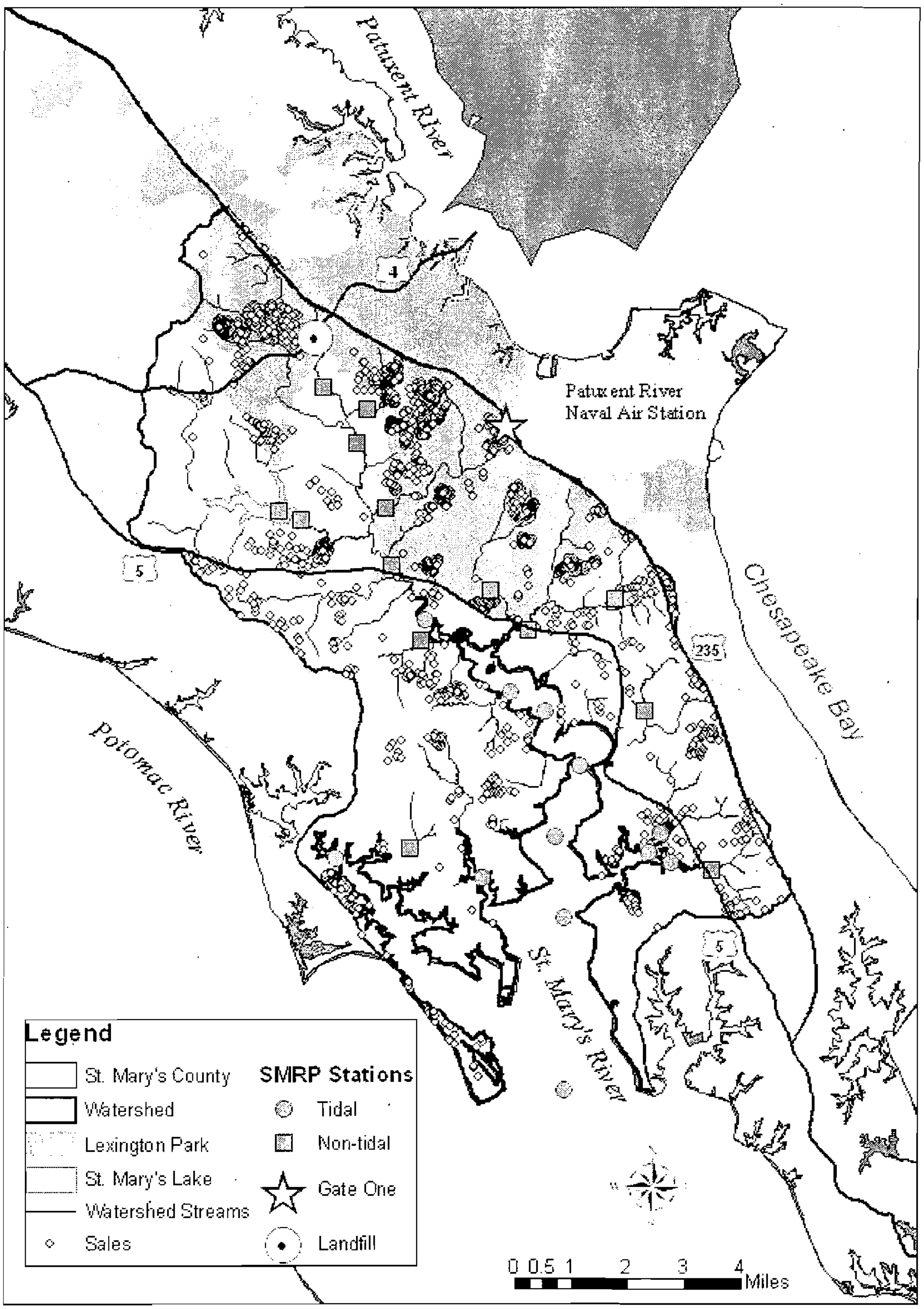


**Figure 2 - The St. Mary's River Watershed**



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APPENDIX A – Qualitative Descriptions of the Index of Biological Integrity and Fish Index of Biological Integrity Scales

**Table 4 – Index of Biological Integrity and Fish Index of Biological Integrity**

| Score Range | Narrative Rating/Description (compared to reference minimally impacted stream) |
|---|---|
| 4.0 – 5.0 | **Good** - minimally impacted |
| 3.0 – 3.9 | **Fair** - some aspects of biological integrity may not resemble the qualities of referenced streams |
| 2.0 – 2.9 | **Poor** - Significant deviation from reference conditions, many aspects not resembling the qualities of referenced, some degradation |
| 1.0 – 1.9 | **Very Poor** - Strong deviation from reference conditions, most aspects not resembling qualities referenced, severe degradation |

Maryland Dept. of Natural Resources Report: "Development of a Benthic Index of Biotic Integrity for Maryland Streams." CBWP-EA-98-3, 1998.

APPENDIX B – Qualitative Descriptions of the Aesthetic Rating Scale

**Table 5 – Aesthetic Rating**

| Score | Narrative Rating/Description |
|---|---|
| 16 - 20 | **Optimal** – little evidence of humans, natural state |
| 11 - 15 | **Sub-Optimal** – minor evidence of humans, minor disturbance of vegetative state |
| 6 - 10 | **Marginal** – Moderate amounts refuse, and vegetative disturbance, channel alteration present |
| 0 - 5 | **Poor** – Abundant, unsightly amounts refuse, nearly complete lack of vegetative, channel alteration extensive |

Maryland Dept. of Natural Resources Report: "Development of a Benthic Index of Biotic Integrity for Maryland Streams." CBWP-EA-98-3, 1998.

# Reductions in the Economic Value of Walleye and Salmon Fishing Due to Low Water Levels at Lake Sakakawea, North Dakota

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## Abstract:

The five foot reduction in water levels at Lake Sakakawea, North Dakota between 2002 and 2003 due to drought conditions and USACE management of the Missouri River, reduced the economic value of recreational fishing to the State by $2.6 million, or 6.6% of pre-drought values. Consumer surplus losses estimated with a travel cost model were approximately twice as large as reduced daily expenditures adjusted for site-substitution effects. Future (2004-2011) reductions in economic values are expected to be $90.2 million, or $11.2 million per year.

## Introduction:

This study calculates economic losses associated with reduced walleye and salmon fishing opportunities at Lake Sakakawea, North Dakota due to U.S. Army Corps of Engineers (USACE) management practices on the Missouri River which have reduced recreational access, impaired water quality, and reduced both catch rates and angler participation. Economic losses are defined as reductions in both daily expenditures and consumer surplus values which required accurate estimates of:

- Catch rates and angler participation over time for salmon and walleye angling.
- Sakakawea-specific daily expenditure and consumer surplus values.
- In-state substitution patterns for Sakakawea anglers.
- Likely future reductions in angler days due to continued low water levels.

This information was then used to calculate actual economic damages associated with reduced walleye and salmon angling at Lake Sakakawea resulting from low water levels in 2003 as well as likely economic damages in the next eight years (2004-2011) if low water levels at Lake Sakakawea continue.

## Background:

Water levels in 2003 at Sakakawea were on average 1825 feet above mean sea level (ft-msl), or 5 feet lower than in 2002 (Power, 2003). This resulted in: lower catch rates (64% for walleye, and 33% for salmon) and reduced angler days 5.6% for resident walleye anglers, 29% for resident salmon anglers, and 35% for non-resident salmon anglers. In contrast, non-resident walleye angler days increased by 6% (Table 1). Not all reduced angler days were lost entirely to the state because 15.2% of resident walleye anglers substitute fishing at Sakakawea with other sites in North Dakota versus 16.5% of resident salmon anglers (who would fish for walleye either at Sakakawea or other places in the state). Similarly, 9.9% of walleye anglers would substitute Sakakawea fishing for other in-state sites, whereas 0% of salmon anglers would fish at other North Dakota locations.

**Reduced Angler Days at Lake Sakakawea (2000-2003)**

The value of baseline (2000-2002), 2003, and future annual expenditure values associated with fishing on Lake Sakakawea are calculated by multiplying estimates of angler days at Lake Sakakawea based on year 2000 and 2003 NDGF creel surveys (Brooks and Hendrickson, 2000 and 2003) by daily variable and fixed expenditure estimates.

The impact of reduced fishing quality from 2000 to 2003 at Lake Sakakawea (due to an average 5 foot lower water level at the lake and reduced catch rates) was a reduction in resident angler days by 10% for all species), 7.1% for walleye, and 28.9% salmon. In contrast, non-resident angler days increased (2.4% for all species and 6% for walleye), but declined 35% for salmon. Modest reductions in resident angler days for walleye are likely a result of the fact that walleye catch rates at Sakakawea, although reduced compared to previous years, are still good compared to many other walleye fisheries in this region of the country.

**Variable Angler Expenditures**

Year 2001 North Dakota Game and Fish (NDGF) expenditure survey data based on mail surveys of resident and non-resident fishing license holders was manipulated to obtain Sakakawea specific daily angler expenditures: Non-Missouri River system anglers, and respondents with expenditures greater than two standard deviations above mean values (outliers) were excluded. Remaining observations were then averaged across residences (zipcodes) and then weighted by observed frequency of visit to Sakakawea (zipcodes of visitors from creel surveys). Resulting daily expenditure values were \$41 for residents and \$84 for non-residents.

**Site Substitution and Net Expenditure Losses to the State**

Estimating actual economic losses to North Dakota associated with reduced angler days as a result of declining fishing quality at Lake Sakakawea requires accounting for the angler days that will be substituted at alternative sites within the state. It is important to estimate how many lost Sakakawea angler days are negated by the generation of additional angler days at other North Dakota locations. We estimated these 'substitution effects' associated with a reduction in fishing quality at Lake Sakakawea by asking 2003 creel survey respondents a series of questions regarding their historical, current, and planned fishing activities at Sakakawea and their intentions to continue fishing at Sakakawea or fish elsewhere with the hypothetical onset of reductions in fishing quality.

If fishing quality at Sakakawea would continue 59% of resident anglers would continue fishing at Sakakawea, 18% said they would fish less at Sakakawea but nowhere else (no substitutes), while 23% would find alternative (substitute) sites. The proportion of anglers stating they would go to substitute sites were 31% for non-residents, 23% for resident walleye anglers, 19% for resident salmon anglers, and 28% for non-resident walleye anglers. The distance anglers traveled to Sakakawea did not appear to influence their propensity to fish at substitute sites. However, anglers living close to alternative fishing sites are more prone to fish at those sites than those living far from any substitute fishing locations. Since 84% of Sakakawea salmon anglers said they would continue to fish at Sakakawea if fishing for other species (i.e. walleye) remained good, yet only 59% said they would continue at Sakakawea if fishing for all species were poor, we conclude that about 25% of Sakakawea salmon anglers would switch to fishing for walleye if the salmon fishery collapsed.

The location of Sakakawea anglers stated substitute sites are summarized in Table 3. Most resident anglers indicated that they would utilize substitute sites that are within North Dakota, while non-resident anglers would utilize substitute sites primarily in Montana and North Dakota. In most cases, Canada was mentioned as the third most common location for substitute fishing sites.

The final assessment of lost Sakakawea angler days to North Dakota due to poor fishing conditions requires multiplying the percentage of anglers using substitute sites by the percentage of substitute sites that are located in North Dakota. The resulting net in-state substitution effects are shown in Table 4, indicated that 15% of lost resident walleye angler days will be substituted for other walleye angler days in the state (and are therefore not a loss to the state as a whole), whereas 16.5% of lost resident salmon angler days are not really lost (because 25% of salmon anglers are assumed to switch to walleye fishing). Only 10% of non-resident walleye angler days are expected to be replaced by fishing at alternative North Dakota sites. This makes sense because many of these non-residents have substitute walleye sites that are much closer to their origin than such as Devils Lake which is the primary substitute walleye fishery in North Dakota. Finally, it is expected that none of the lost angler days associated with non-resident salmon anglers will be substituted for in North Dakota because of the lack of alternative salmon fisheries in the state.

Combining observed reductions in angler days at Lake Sakakawea between 2000 and 2003 due to low water levels and reduced fishing quality with net in-state substitution effects enables us to quantify how many of the reduced Sakakawea angler days were actually lost to the state. A total of 16,925 resident walleye anglers and 9,125 resident salmon anglers were lost to the state due to low water quality and declining fishing quality at Lake Sakakawea in conjunction with a lack of suitable substitute sites between 2000 and 2003. Lost non-resident salmon days were 898 while non-resident walleye angler days increased by 1,799

**Travel Cost Modeling of Consumer Surplus Values For Walleye and Salmon Fishing**

To estimate the consumer surplus values associated with Lake Sakakawea angling, a travel cost model was estimated using demand analysis questions in conjunction with the 2003 creel survey of anglers on Lake Sakakawea. Respondents were about their fishing trips and days per trip to Lake Sakakawea in 2003. Information regarding their income and education levels, along with distance to a substitute fishing sites was solicited. The economic survey data when combined with the 2003 creel survey data provided enough information to estimate an individual travel cost model for Lake Sakakawea. The travel cost model was specified as:

TRIPS = $\beta_0$ + $\beta_1$(TRAVEL COST) + $\beta_2$(PRICE OF SUBSTITUTE SITE) + $\beta_3$(WALLEYE CATCH RATE) + $\beta_4$(INCOME) + $\beta_5$(PARTY SIZE) + $\beta_6$(DAYS PER TRIP) + $\beta_7$(NORTH DAKOTA RESIDENT*TRAVEL COST).

The $\beta$'s are the parameters of the model to be estimated and represent the marginal change in trips associated with a one unit change in the parameter's variable. TRIPS is the observed number of trips taken to Lake Sakakawea. TRAVEL COST serves as the proxy for the price of taking a trip to fish at Lake Sakakawea and is composed of mileage costs, travel time costs and other variable out-of-pocket expenses for fishing at Lake Sakakawea. Mileage cost was estimated as the variable costs of gasoline, oil and tire wear assumed to be $0.12 per mile (2003 AAA rate) times roundtrip miles from angler's residence to Lake Sakakawea. Travel time cost was estimated using 30% of the angler's wage rate per hour (calculated from their income) times roundtrip miles/60 miles per hour. Other variable costs were estimated from the 2001 Fishing Expenditure Survey discussed elsewhere in this report.

The PRICE OF SUBSTITUTE SITE was calculated using the same formulas as Travel Cost with the exception of roundtrip miles from residence to Lake Sakakawea. Substitute sites were either stated by the respondent or inferred based on the location of the angler's residence. The WALLEYE CATCH RATE served as a proxy for fishing quality at the lake. It was calculated from the 2003 creel survey data by dividing total walleye caught per party by total hours fished per party. Therefore, WALLEYE CATCH RATE measures the number of walleye caught per hour. A salmon catch rate could not be calculated given the lack of salmon anglers in the 2003 creel survey.

INCOME is the annual household gross income of the angler. PARTY measures party size in number of anglers in the respondents fishing group. DAYS measures the days per trip of each angler surveyed. NORTH DAKOTA RESIDENT*TRAVEL COST measures the incremental difference in travel costs for residents of North Dakota. Summary statistics for the variables included in the estimated model are provided in Table 5.

The travel cost model was estimated on 1,622 observations using both Ordinary Least Squares and Negative Binomial estimators. Ordinary Least Squares is an appropriate estimator when the dependent variable (TRIPS) is continuously defined and measured. However, trips are reported and taken in whole integers only. Therefore, an appropriate estimator is the Negative Binomial[12]. In addition, the predicted number of trips taken by our sample was estimated to be 11,107 by the OLS model and 19,070 by the negative binomial model; in fact, the actual number of trips taken by our sample was 19,292. Therefore, the negative binomial model fits the data better than the OLS model. Results of the negative binomial model are reported in Table 6

Most of the coefficient estimates are statistically different than zero at a 95% confidence level or greater. The price coefficient (TRAVEL COST) is significant and negative meaning we have estimated a demand curve based on the behaviorally observed travel cost data. For every dollar increase in price, trips demanded decline. PRICE OF SUBSTITUTE SITE is significant and positive, showing that the higher the cost of a substitute will result in more trips that will be taken to Lake Sakakawea. The WALLEYE CATCH RATE shows trip behavior is directly related to fishing quality as measured by walleye catch rate. NORTH DAKOTA RESIDENT*TRAVEL COST shows North Dakota anglers' travel costs to Lake Sakakawea are statistically lower than non-residents' travel costs.

The area below the estimated demand curve but above the price line defines the level of consumer surplus. Consumer surplus (CS) is calculated based on the formula for a semi-log (natural-log of TRIPS) functional form, or CS/Trip = $1/-(\beta_1)$ for non-residents and CS/Trip = $1/-(\beta_1 + \beta_7)$ for residents of North Dakota. This amounts to \$149.68 per trip for non-resident anglers at Lake Sakakawea in 2003 and \$178.88 per trip for resident anglers at Lake Sakakawea

[12] The negative binomial is a count data estimator in the compound Poisson family. The negative binomial is less restrictive than the Poisson by not forcing equality between the mean and the variance of the dependent variable. When the mean and the variance of the dependent variable are not equal, the Poisson model suffers from over dispersion. In our data, the mean and variance of TRIPS are not equal. Therefore, the negative binomial estimator was used.

in 2003. At an average of 2.27 days per trip, non-residents' consumer surplus is $66.02 per day and residents' consumer surplus is $78.91 per day. Consumer surplus for non-resident anglers is lower than residents' because of the higher price (distance-related) travel to Sakakawea.

**Conclusion: Economic Impacts of Low Water Levels**

The total annual economic value of walleye and salmon fishing at Lake Sakakawea from 2000 to 2002 (prior to the 2003 low water season) is estimated at $39.2 million ($15.7 million from daily and expenditures and $23.6 million from consumer surplus values. Again, these these estimates are based on adjusted NDGF license survey variable expenditures, a multiplier effect of 1.13 applied to variable expenditures, and consumer surplus values estimated via a travel cost model.

Economic losses from 2002 to 2003 associated with a 5 foot drop in the elevation of Lake Sakakawea and reduced walleye and salmon fishing quality were $2,609,138 ($932,579 from reduced daily expenditures and $1,676,559 in lost consumer surplus values). This corresponds to a loss of 6.6% of the annual economic value of walleye and salmon fishing at Sakakawea in 2003 (prior to critically low water levels).

Based on historical observations concerning water levels, fishing quality and angler participation at Lake Sakakawea (from creel survey data that includes the 1987 to 1992 drought period), we estimate total economic losses at Lake Sakakawea to be in excess of $90 million over the next eight years due to low water levels and reductions in walleye and salmon angling days. This includes about $33 million in lost variable expenditures and $57 million in lost consumer surplus values. Economic losses associated with lost resident angler days total $81 million versus $9 million in economic losses associated with non-resident anglers (Table 8).

**References:**

Brooks, L. and J. Hendrickson. 2000. Angler Use and Sport Fishing Catch Survey on Lake Sakakawea, North Dakota, May 1 Through September 30, 1994. North Dakota Game and Fish Department Report F-2-R-41-3-2.

Brooks, L. and J. Hendrickson. 2003. Angler Use and Sport Fishing Catch Survey on Lake Sakakawea, North Dakota, May 1 through October 24, 2003. North Dakota Game and Fish Department Report.

Power, Greg. 2003. Memorandum on Lake Sakakawea Fishery Information. March 13, 2003. NDGF Department.

**Table 1. Differences in Water Levels, Fishing Quality, and Angler Effort and Origin: Lake Sakakawea, 2000 to 2003**

| | 2000 | 2003 | Difference |
|---|---|---|---|
| Lake Elevation (avg. summer, ft-msl) | 1830 | 1825 | - 5 feet |
| Total Angler Days [1] | | | |
| -All species | 309,754 | 282,937 | - 8.7 % |
| -Walleye [2] | 269,468 | 254,360 | - 5.6 % |
| -Salmon | 34,073 | 24,050 | - 29.0 % |
| | | | |
| Fish Caught Per Angler Hour | | | |
| - Walleye | 0.635 | 0.228 | - 64 % |
| - Salmon | 0.015 | 0.010 | - 33% |
| Angler Reported Quality of Fishing [3] | | | |
| - Walleye anglers | 3.1 | 2.5 | - 19 % |
| - Salmon anglers | 3.5 | 2.1 | - 40% |

*1. Angler days are based on NDGF creel survey reports (Brooks and Hendrickson, 2000, and 2003) plus 15% to account for angling during non-creel survey periods while other statistics are based on actual creel survey data.*
*2. Walleye statistics do not include sauger*
*3. Quality: Excellent (5), Good (4), Average (3), Fair (2), Poor (1)*

**Table 2 Intentions of All Anglers if All Fishing Quality Declines**

| | Continue Fishing the Same Amount at Sakakawea | Fish Less at Sakakawea and Nowhere Else (No Substitutes) | Fish Elsewhere |
|---|---|---|---|
| Resident Anglers | 59% | 18% | 23% |
| Non-Resident Anglers | 54% | 15% | 31% |
| Resident Walleye Anglers | 59% | 18% | 23% |
| Resident Salmon Anglers | 69% | 13% | 19% |
| Non-Resident Walleye Anglers | 52% | 20% | 28% |
| Non-Resident Salmon Anglers | NA | NA | NA |
| Anglers within 50 miles | 50% | 26% | 24% |
| Anglers within 50-99 miles | 58% | 18% | 24% |
| Anglers within 100-299 miles | 53% | 22% | 26% |
| Anglers farther than 300 miles | 60% | 16% | 24% |

Table 3. Location of Substitute Sites Due to a Decline in Sakakawea Fishing Quality

| | N.D. | S.D. | MN | MT | Other States | Canada |
|---|---|---|---|---|---|---|
| Resident Anglers | 73% | 5% | 2% | 3% | 0% | 7% |
| Non-Resident Anglers | 33% | 15% | 5% | 38% | 2% | 7% |
| Anglers traveling <50 miles | 60% | 3% | 5% | 0% | 0% | 15% |
| Anglers traveling 50-99 miles | 73% | 3% | 1% | 3% | 0% | 2% |
| Anglers traveling 100-299 miles | 60% | 10% | 1% | 22% | 0% | 4% |
| Anglers traveling >300 miles | 59% | 18% | 7% | 7% | 2% | 5% |

**Table 4. Substitution Effects Among Lake Sakakawea Anglers**

| | Anglers Using Substitute Sites With Declining Fishing Quality | Percentage of Substitute Sites in ND | Net In-State Substitution Effects |
|---|---|---|---|
| Resident Walleye Anglers | 23% | 66% | 15.2% |
| Resident Salmon Anglers | 25%[1] | 66%[1] | 16.5% |
| Non-Resident Walleye Anglers | 30% | 33% | 9.9% |
| Non-Resident Salmon Anglers | 100% | 0%[2] | 0% |

*1. Assumed to switch to walleye fishing*
*2. Estimated because no other salmon fishing sites exist in North Dakota*

**Table 5 Travel Cost Model: Summary Statistics (n = 1,622)**

| VARIABLE | MEAN | STD.DEV. | MIN | MAX |
|---|---|---|---|---|
| TRIPS | 11.87 | 11.89 | 1 | 90 |
| TRAVEL COST (price per trip) | $95.14 | 118.53 | $0.74 | $1624.28 |
| PRICE OF SUBSTITUTE SITE | $112.78 | 116.55 | $6.89 | $1667.53 |
| WALLEYE CATCH RATE | 0.50 | 0.71 | 0 | 10 |
| INCOME (1.000s) | $65.81 | 21.67 | 12.5 | 225 |
| PARTY SIZE | 2.41 | 0.91 | 1 | 7 |
| DAYS | 2.27 | 1.55 | 1 | 15 |
| NORTH DAKOTA RESIDENT*TRAVEL COST | 57.10 | 56.67 | 0 | 589.70 |

**Table 6 Travel Cost Demand Model, Negative Binomial (n = 1,622)**

| VARIABLE | COEFFICIENT | STANDARD ERROR |
|---|---|---|
| Constant | 2.7174** | 0.1128 |
| TRAVEL COST | -0.0067** | 0.0007 |
| PRICE OF SUBSTITUTE SITE | 0.0048** | 0.0007 |
| WALLEYE CATCH RATE | 0.1194** | 0.0311 |
| INCOME (1,000s) | 0.0005 | 0.0012 |
| PARTY SIZE | 0.0244 | 0.0311 |
| DAYS | -0.2162** | 0.0228 |
| NORTH DAKOTA RESIDENT* TRAVEL COST | 0.0011* | 0.0005 |
| Alpha (over dispersion constant) | 0.6544** | 0.0266 |
| | | |
| Log-Likelihood | -5482.14 | |

*Coefficient is statistically different than zero at a 95% confidence level; **confidence at the 99% level.

**Table 7 Angler Days Lost to North Dakota Due to Reduced Sakakawea Fishing (2000-2003)**

| | Change in Angler Days at Sakakawea (2000-2003)[1] | Reduced Angler Days Substituted with In-State Sites[2] | Net Loss/Gain of Angler Days to the State |
|---|---|---|---|
| Resident Walleye | - 16,925 | 2,569 | -14,356 |
| Resident Salmon | - 9,125 | 1,506 | - 7,619 |
| Non-Resident Walleye | + 1,799 | NA | + 1,799 |
| Non-Resident Salmon | - 898 | 0 | - 898 |

*1. From 2000 and 2003 creel surveys* *2. Net substitution effects plus % anglers without any substitutes*

**Table 8. Projected Economic Losses Due to Changes at Lake Sakakawea, 2004-2011.**

| Measure | Resident Walleye Fishing | Resident Salmon Fishing | Non-Resident Walleye Fishing | Non-Resident Salmon Fishing |
|---|---|---|---|---|
| Cumulative Predicted Angling Days | 1,349,231 | 17,995 | 194,436 | 1,061 |
| Cumulative Potential Angling Days | 1,914,096 | 252,272 | 241,744 | 20,312 |
| Gross Lost Angler Days without Substitution Effects | -564,865 | -234,277 | -47,308 | -19,251 |
| Net Lost Angler Days with Substitution Effects | -479,005 | -195,622 | -42,624 | -19,251 |
| Daily Variable Expenditures | $41 | $41 | $84 | $84 |
| Consumer Surplus per Day | $79 | $79 | $66 | $66 |
| Total Lost Daily Variable Expenditures | -$19,639,219 | -$8,020,484 | -$3,580,456 | -$1,617,117 |
| Total Lost Consumer Surplus | -$37,841,423 | -$15,454,103 | -$2,813,215 | -$1,270,592 |
| Total Economic Losses | - $57,480,642 | -$23,474,587 | -$6,393,671 | -$2,887,709 |
| Grand Total Economic Loss | - $90,236,609 | | | |

# Empirical Strategies for Incorporating Weak Complementarity into Continuous Demand System Models

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First Draft: February 24, 2004
Current Draft: May 24, 2004

I thank Matt Massey for generously providing the data used in the empirical application. Dan Phaneuf, Michael Hanemann, Kerry Smith, and seminar participants at Stanford, New Hampshire, and Nevada also provided helpful comments. All remaining errors are my own.

**Empirical Strategies for Incorporating Weak Complementarity into Continuous Demand System Models**

## Abstract.

This paper conceptually and empirically compares alternative strategies for incorporating weak complementary into continuous demand system models. The repackaging, integrating back, and discontinuity approaches are evaluated in terms of their behavioral implications and potential usefulness for applied research. The conceptual review suggests that the repackaging approach offers the most flexibility and tractability. The empirical comparison suggests that qualitatively similar policy inference arises from previously employed repackaging approaches. These estimates can be similar to use-related welfare estimates derived from non-weakly complementary models using a decomposition approach suggested by Herriges, Kling, and Phaneuf (2004), although the latter are sensitive to arbitrary assumptions about how to decompose use and nonuse values.

**JEL Classification:** D120, Q260, C240, C110

## 1. Introduction

In applied demand analysis, neoclassical consumer theory provides guidance for structuring relationships among quantities, prices, and income. The theory is, however, noticeably silent with respect to how the quality attributes of goods should enter these relationships. As a result the analyst has considerable discretion when introducing goods' quality attributes into consumer demand models. Because a significant and growing number of measurement questions arise in the context of quality change, this reality challenges the researcher to develop preference and demand specifications that defensibly incorporate goods' quality attributes.

When developing these specifications, the analyst can sometimes statistically discriminate between alternative hypotheses about how quality enters preference and demand relationships. Otherwise, intuition is the analyst's only guide. One untestable but intuitive restriction on how quality attributes enter these relationships that Mäler (1974) and Bradford and Hildebrandt (1977) proposed three decades ago is weak complementarity. When a good and its quality attributes are weak complements, the individual only values marginal improvements in the good's quality attributes if she consumes it. This restriction implies that all value derived from changes in a good's quality attributes arises through consumption. Whether implicit or explicit, this restriction has been incorporated into preferences for virtually every valuation exercise that relies exclusively on revealed preference data.

This paper conceptually and empirically compares alternative strategies for incorporating weak complementarity into continuous demand system models. Three different strategies – the repackaging approach (Fisher and Shell (1968)), the integrating back approach (Larson (1991)), and the discontinuity approach (Bockstael, Hanemann, and Strand (1986)) – are evaluated in terms of their behavioral implications and potential empirical usefulness. One of the main implications from a conceptual comparison of the approaches is that only the repackaging approach is likely to offer applied researchers much guidance and flexibility when developing weakly complementary demand specifications. An empirical comparison across three weakly complementary specifications developed within the repackaging framework is conducted with a beach recreation data set and estimated within the Bayesian statistical framework. The main empirical finding is that welfare estimates for the loss of beach width are qualitatively similar across the alternative repackaging specifications considered. These estimates can be similar to use-related welfare estimates derived from non-weakly complementary models using a decomposition approach suggested by Herriges, Kling, and Phaneuf (2004) (hereafter HKP), although the latter are sensitive to arbitrary judgments about how to decompose use and nonuse values.

The remainder of the paper is structured as follows. For perspective, section 2 reviews how economists have proposed introducing goods' quality attributes explicitly into continuous demand system models with a special emphasis on the role of weak complementarity. Section 3 then critically reviews the repackaging, integrating back, and discontinuity approaches to developing weakly complementary demand models. Section 4 summarizes the specifications to be compared in the empirical application, and Section 5 briefly summarizes the 1997 mid-Atlantic beach recreation data used in the comparison. Section 6 follows with the parameter estimates for the alternative models, and section 7 discusses the welfare scenario, estimation strategy, and empirical results. Section 8 concludes with a discussion of the implications of the paper's findings for future research.

## 2. Introducing Quality into Demand Systems & the Role of Weak Complementary: A Review

As discussed in Hanemann (1982), economists have historically exploited one of two generic frameworks for explicitly incorporating the quality attributes of heterogeneous goods into demand system models. The first approach was pioneered by Houthakker (1952-3) and Theil (1952-3) and assumes that goods can be grouped into categories or classes based on their similar functions and characteristics. Within each class, the individual is assumed to consume at most one good, and the similar goods collectively form a perfect continuum of alternatives over the relevant support of all quality attributes. Consumer preferences in this setup can be represented by a direct utility function, $U(\boldsymbol{x},\boldsymbol{Q},z)$, where $\boldsymbol{x}$ corresponds to an $N$ dimensional vector of consumption quantities for the different classes of goods, $\boldsymbol{Q}=[\boldsymbol{q}_1,\ldots,\boldsymbol{q}_N]$ is a matrix of quality attributes for all $N$ good classes, and $z$ is an essential Hicksian composite good. The price of consuming a particular type of good in the $j$th class is $p_j(\boldsymbol{q}_j)$, where $p_j(\cdot)$ is a smooth, continuous "hedonic" price function.[13] Together these assumptions imply that the consumer's problem can be succinctly stated as:

$$(1) \qquad \max_{x,Q}\{U(\boldsymbol{x},\boldsymbol{Q},z)\} \quad \text{s.t. } \sum_j p_j(\boldsymbol{q}_j)x_j+z=y,\ \boldsymbol{x}\geq 0,\ \&\ \underline{\boldsymbol{q}}_j\leq\boldsymbol{q}_j\leq\overline{\boldsymbol{q}}_j,\forall j,$$

[13] Alternatively, Theil (1956) interprets $p_j(\boldsymbol{q}_j)$ as a quality adjusted price index for the $j$th commodity group, which in turn implies that $x_j$ is a quality adjusted quantity index.

where $\underline{\boldsymbol{q}}_j$ and $\overline{\boldsymbol{q}}_j$ are the upper and lower bounds of the support for $j$th good's quality attributes. Equation (1) can be interpreted as a multivariate generalization of Rosen's (1974) discrete-continuous formulation of the consumer's problem when $\boldsymbol{x}$ is a scalar equal to one. Two distinguishing features of this representation of the consumer's problem are worth emphasizing. Both the quantities and quality attributes of goods enter as endogenous arguments of the individual's preference ordering that imply first order conditions that implicitly define the consumer's optimal consumption bundle. In addition, the budget constraint may be highly nonlinear in quality attributes, and thus some structure must be placed on the hedonic price functions for unique solutions to exist.[14]

A second strategy for explicitly incorporating quality attributes into demand system models may be preferable when the continuity assumptions that underlie Houthakker and Theil's approach are inappropriate. As Lancaster (1966) and Mäler (1974) argue, the quality attributes of a finite set of goods can be thought of as exogenous fixed factors or rationed goods. Thus quality attributes in this formulation are nothing more than additional parameters that enter the individual's preference ordering. A good's price may also depend on its quality attributes, but because quality attributes are exogenous to the individual, the dependence need not be made explicit. The consumer's problem in this setting can be stated as:

$$(2) \qquad \max_{x}\left\{U(\boldsymbol{x},\boldsymbol{Q},z)\right\} \quad \text{s.t. } \boldsymbol{p}^{\mathsf{T}}\boldsymbol{x}+z=y,\boldsymbol{x}\geq 0,$$

where $\boldsymbol{x}$ contains all quality differentiated goods. Notice in (2) that only $\boldsymbol{x}$ is endogenous and the budget constraint is linear. In some sense, this structure is the natural extension to the traditional formulation of the consumer's problem where the utility function's dependence on the quality

[14] Rosen (1974) discusses how market forces will result in the hedonic price functions embodying this additional structure.

attributes of goods is implicit. Because the rationed goods approach does not rely on continuity assumptions that may be implausible in many applications, it is often the preferred approach for incorporating quality attributes explicitly except in the extreme but not altogether uncommon situations considered by Rosen. Thus in the remainder of the paper the discussion is couched exclusively in terms of the rationed goods approach.

When quality is introduced explicitly into consumer demand models, a natural question arises: does neoclassical consumer theory suggest any structure for how quality enters preference and demand functions? In general, the answer is no. The restrictions on preference and demand relationships implied by traditional consumer theory represent the minimal set of assumptions necessary to guarantee a solution to the consumer's constrained optimization problem. So long as preferences satisfy these conditions and the consumer's affordable budget set is convex, a unique solution to the consumer's problem is guaranteed regardless of how goods' quality attributes enter preferences. Thus although consumer theory has much to say about the relationships among quantities, prices, and income in preference and demand functions, it has nothing to say about how quality attributes enter these relationships. This reality is in some sense liberating to the analyst, but it also place a significant burden on her to develop defensible empirical specifications that explicitly incorporate goods' quality attributes.

When developing these specifications, intuition is often the analyst's only guide, although statistical criteria can sometimes be used to discriminate among competing hypotheses on how quality attributes enter demand and preference functions. Weak complementarity, which is the focus of this paper, represents an intuitive but untestable restriction. In essence, weak complementarity implies that all value derived from the quality attributes of a good arise exclusively through the good's use. Two conditions must be satisfied for the preference restriction to hold: 1) the good must be nonessential; and 2) if the good is not consumed, the individual does not benefit from marginal improvements in its quality attributes, i.e., $\partial U(\boldsymbol{x},\boldsymbol{Q},z)/\partial \boldsymbol{q}_j = 0$ if $x_j = 0, \forall j$. Weak complementarity is not a testable restriction because the analyst cannot distinguish between a utility function $U(\boldsymbol{x},\boldsymbol{Q},z)$ that satisfies conditions 1) and 2) and a monotonic transformation of $U(\boldsymbol{x},\boldsymbol{Q},z)$, say $u(U(\boldsymbol{x},\boldsymbol{Q},z),\boldsymbol{Q})$, that does not with just revealed preference data for $\boldsymbol{x}$.[15,16]

Following Smith and Banzhaf (2004), weak complementarity's implications for the structure of preferences are represented graphically in Figure 1 in the context of a simple two good ($x$,$z$) model. The horizontal axis measures the quantity of $x$ consumed while the vertical axis measures the consumption level of $z$. Three indifference curves are drawn that correspond to the same level of utility (i.e., $U(\boldsymbol{q}^0) = U(\boldsymbol{q}^1) = U(\boldsymbol{q}^2)$) but different levels of quality ($\boldsymbol{q}^0 > \boldsymbol{q}^1 > \boldsymbol{q}^2$) where utility is assumed to be strictly increasing in quality. The first condition of weak complementarity requires that these indifference curves must intersect the vertical axis, and the second that they intersect the vertical axis at the same point (in figure 1, point A). Assuming preferences are continuous, they also exhibit what Smith and Banzhaf refer to as the "fanning" property – as $x$ increases, the distance between the indifference also increases.

In the context of a single quality differentiated good, Mäler (1974) showed how weak complementarity and Hicksian demand functions can be used to construct theoretically

[15] This point should not be interpreted as suggesting that weak complementarity implies preference functions are sensitive to all monotonic transformations. Although weak complementarity does rule out monotonic transformations of $U(\boldsymbol{x},\boldsymbol{Q},z)$ that change the marginal rates of substitution among $\boldsymbol{x}$, $\boldsymbol{Q}$, and $z$, it does not rule out those that do not.

[16] However, if both revealed and stated preference data are present, it may be possible to test whether weak complementarity is a valid assumption.

consistent welfare measures. A key concept in his derivation is the Hicksian "choke" price, or the minimal price that drives the consumer's Hicksian demand for $x$ to zero. Hicksian choke prices will in general depend on $\boldsymbol{q}$ and $\bar{U}$, the relevant utility level. Graphically in Figure 1, they correspond to the slopes of the indifference curves $U(\boldsymbol{q}^0)$, $U(\boldsymbol{q}^1)$, and $U(\boldsymbol{q}^2)$ evaluated at $x = 0$. Mäler demonstrated that weak complementarity implies that the difference between the compensating variations associated with a price change from the observed to the Hicksian choke price evaluated separately at $\boldsymbol{q}^0$ and $\boldsymbol{q}^1$ represents the Hicksian consumer surplus, $CS^H(\boldsymbol{q}^0,\boldsymbol{q}^1)$, arising from the quality change. This can be seen below in equation (3):

$$
\begin{aligned}
CS^H(\boldsymbol{q}^0,\boldsymbol{q}^1) &= E(\bar{p},\boldsymbol{q}^0,\bar{U}^0) - E(\bar{p},\boldsymbol{q}^1,\bar{U}^0)\\
&= E(\bar{p},\boldsymbol{q}^0,\bar{U}^0) - E(\bar{p},\boldsymbol{q}^1,\bar{U}^0) + E(\hat{p}(\boldsymbol{q}^1,\bar{U}^0),\boldsymbol{q}^1,\bar{U}^0) - E(\hat{p}(\boldsymbol{q}^0,\bar{U}^0),\boldsymbol{q}^0,\bar{U}^0)\\
&= \left[E(\hat{p}(\boldsymbol{q}^1,\bar{U}^0),\boldsymbol{q}^1,\bar{U}^0) - E(\bar{p},\boldsymbol{q}^1,\bar{U}^0)\right] - \left[E(\hat{p}(\boldsymbol{q}^0,\bar{U}^0),\boldsymbol{q}^0,\bar{U}^0) - E(\bar{p},\boldsymbol{q}^0,\bar{U}^0)\right],\\
&= \int_{\bar{p}}^{\hat{p}(q^1,\bar{U}^0)} x^h(p,\boldsymbol{q}^1,\bar{U}^0)dp - \int_{\bar{p}}^{\hat{p}(q^0,\bar{U}^0)} x^h(p,\boldsymbol{q}^0,\bar{U}^0)dp\\
&= CV(\bar{p},\hat{p}(\boldsymbol{q}^1,\bar{U}^0),\boldsymbol{q}^1,\bar{U}^0) - CV(\bar{p},\hat{p}(\boldsymbol{q}^0,\bar{U}^0),\boldsymbol{q}^0,\bar{U}^0)
\end{aligned}
\tag{3}
$$

where $\bar{p}$ is the observed market price, $\hat{p}(\cdot,\bar{U}^0)$ is the Hicksian choke price, and $E(\cdot,\cdot,\bar{U}^0)$ is the expenditure function. The first line of equation (3) is simply the definition of the Hicksian consumer surplus associated with the quality change from $\boldsymbol{q}^0$ and $\boldsymbol{q}^1$, the second adds in the expression $E(\hat{p}(\boldsymbol{q}^1,\bar{U}^0),\boldsymbol{q}^1,\bar{U}^0) - E(\hat{p}(\boldsymbol{q}^0,\bar{U}^0),\boldsymbol{q}^0,\bar{U}^0)$ which equals zero if weak complementarity holds, the third line simply reorganizes terms, and the fourth and fifth exploit Shephard's lemma and the definitions of Hicksian demand functions, $x^h(\cdot)$, and compensating variation, $CV(\cdot)$, respectively.

One can also graphically illustrate Mäler's result in Figure 1. Imagine that relative prices are such that initially the individual's optimal consumption bundle corresponds to the tangency between line 1 and $U(\boldsymbol{q}^0)$. If the price of $z$ is normalized to one, the expenditures necessary to purchase this bundle correspond to the distance between the origin and point B. The compensating variation associated with the price change from baseline prices $\bar{p}$ to the Hicksian choke price $\hat{p}(\boldsymbol{q}^0,\bar{U}^0)$ (i.e., the slope of $U(\boldsymbol{q}^0)$ evaluated at $x = 0$), corresponds to the distance between points B and A. Weak complementarity implies that a decrease in quality from $\boldsymbol{q}^0$ to $\boldsymbol{q}^1$, although it lowers the choke price from $\hat{p}(\boldsymbol{q}^0,\bar{U}^0)$ to $\hat{p}(\boldsymbol{q}^1,\bar{U}^0)$, does not alter the minimum expenditure necessary to achieve $\bar{U}^0$, i.e., the distance between the origins to point A. A decrease in price from $\hat{p}(\boldsymbol{q}^1,\bar{U}^0)$ to $\bar{p}$ results in the individual's optimal consumption bundle adjusting to the point where Line 2 and $U(\boldsymbol{q}^1)$ are tangent. The compensating variation associated with this price change corresponds to the vertical distance between points A and C. Thus when weak complementarity holds, the consumer's Hicksian consumer surplus associated with the degradation in quality (i.e., the vertical distance between points A and C) exactly equals the difference in compensating variations between $\bar{p}$ and $\hat{p}(\boldsymbol{q}^0,\bar{U}^0)$ $\bar{p}$ and $\hat{p}(\boldsymbol{q}^1,\bar{U}^0)$, respectively.

Although Mäler's duality result is elegant and potentially useful to applied researchers when preferences are quasilinear, its practical value is questionable when income effects are present and observable Marshallian and latent Hicksian demands diverge (see Bockstael and McConnell (1993), Palmquist (2004), and Smith and Banzhaf (2004) for discussions). More importantly, the development of virtually all modern empirical demand system models begins with an explicit specification of preferences represented by a direct or indirect utility function. Palmquist (2004) correctly points out that Mäler's motivation for imposing weak complementarity is less relevant in these situations because the analyst knows (or more precisely, assumes) the complete structure of preferences and welfare measurement is conceptually straightforward. Thus Mäler's rationale for imposing weak complementarity has dubious practical value for the current practice of applied demand analysis.[17]

This point does not diminish the intuitive appeal of the assumption in many applied situations, however. The main implication of weak complementarity is that all value associated with a change in a good's quality attributes arise exclusively through its use. In situations where nonuse values are thought to be absent, imposing weak complementarity *a priori* makes good sense. Moreover, in situations where the analyst believes that nonuse values are likely present but not reliably measurable, imposing weak complementarity still may be defensible. As HKP's and this paper's empirical results suggest, welfare measures derived from demand system models that are not consistent with weak complementarity can be significantly different than estimates derived from models that do. These empirical findings might suggest that nonuse values are substantial (Larson (1993)), but such inference would at best be speculative without additional stated preference data that would allow the analyst to identify more reliably total value. Moreover, the total value estimates are arbitrary in the sense that they are conditional on a specific non-weakly complementary preference ordering when in principle there are an infinite number of non-weakly complementary preference orderings that generate the same observable demand functions.

To address this limitation with welfare estimates derived from non-weakly complementary demand models, HKP have proposed decomposing total value into use and nonuse components, disregarding the unreliably measured nonuse component, and reporting only

[17] Of course its relevance could increase in the future with methodological innovations in applied demand analysis.

the use component. Although intuitively sensible in principle, such decomposition approaches are plagued by at least two problems in practice. Similar to total value, the use component of total value will in general depend on the non-weakly complementary preference structure arbitrarily chosen by the analyst. It is straightforward to show that this dependence can only be broken if the analyst restrictively assumes that preferences are quasilinear. Moreover, for policies affecting only a subset of the quality differentiated goods, Flores (2004) has argued that the decomposition will depend in general on how the analyst defines the nonuse value. To understand his argument, imagine a situation where there are two quality differentiated goods but the policy scenario considered only affects the first good's quality attributes. When defining the nonuse component of total value associated with this quality change, obviously the demand for the first good should be restricted to zero before and after the quality change, but depending on how the second good's demand is treated, the nonuse component of value will change unless the cross-price Hicksian demand elasticity between the two is zero. In other words, the nonuse (and in turn the use) component of total value will depend on whether the demand for the first good or the demands for both goods are held at zero before and after the quality change. Combined, these factors suggest that use-related welfare measures derived from decomposition approaches are sensitive to arbitrary assumptions. Of course welfare measures derived from weakly complementary models that *a priori* rule out nonuse values are also arbitrary. The point of the above discussion, however, is to suggest that the analyst might nevertheless want to assume weak complementary when nonuse values are present but not reliably measurable to avoid the difficulties inherent with decomposition approaches.

## 3. Empirical Strategies for Incorporating Weak Complementarity

If the analyst concludes that incorporating weak complementarity into preferences is appropriate, a relevant question is whether there are generic strategies that can be used to guide the development of weakly complementary specifications. In this section three general strategies that have been identified in the valuation literature are discussed – the repackaging, integrating back, and discontinuity approaches. In principle applied researchers can exploit each of the approaches to develop weakly complementary demand systems, but here it is argued that the repackaging approach is likely to be the most useful in practice.

### *3.1. The Repackaging Approach*

Perhaps the oldest and most widely used strategy for developing empirical demand models consistent with weak complementarity is the repackaging approach. Preferences in this framework can be nested within the following general class of direct utility functions:

$$(4)\qquad U\left(f_{11}(x_1,\boldsymbol{q}_1),\ldots f_{1M_1}(x_1,\boldsymbol{q}_1), f_{22}(x_2,\boldsymbol{q}_2),\ldots f_{2M_2}(x_2,\boldsymbol{q}_2),\ldots, f_{N1}(x_N,\boldsymbol{q}_N),\ldots, f_{NM_N}(x_N,\boldsymbol{q}_N), \boldsymbol{x}, z\right),$$

where $f_{ij}(x_i,\boldsymbol{q}_i), \forall i,j$, are alternative subfunctions that share the property that:

$$(5)\qquad \frac{\partial f_{ij}(x_i,\boldsymbol{q}_i)}{\partial \boldsymbol{q}_i} = 0 \text{ if } x_i = 0, \forall i,j .$$

The structure of equation (4) suggests that the $f_{ij}(x_i,\boldsymbol{q}_i)$ subfunctions aggregate or "repackage" $x_i$ and $\boldsymbol{q}_i$ into $M_i$ composite goods from which the consumer ultimately derives utility. It is only through these subfunctions that $\boldsymbol{q}_i$ enters consumer preferences.

There are at least three empirical repackaging specifications that have been utilized in valuation studies. The oldest and most popular is the pure repackaging framework introduced by

Fisher and Shell (1968). In this framework, the primal representation of consumer preferences can be nested within the following class of direct utility functions:

(6) $$U\left(\phi_1(\boldsymbol{q}_1)x_1, \phi_2(\boldsymbol{q}_2)x_2, \ldots, \phi_N(\boldsymbol{q}_N)x_N, z\right),$$

where $\phi_i(\boldsymbol{q}_i) > 0, i = 1, \ldots, N,$ are commonly referred to as pure repackaging parameters. Muellbauer (1976) shows that the implied indirect utility and Marshallian demand functions consistent with (6) respectively take the form

(7) $$V\left(\frac{p_1}{\phi_1(\boldsymbol{q}_1)}, \ldots, \frac{p_N}{\phi_N(\boldsymbol{q}_N)}, y\right),$$

(8) $$x_i = \frac{1}{\phi_i(\boldsymbol{q}_i)} g_i\left(\frac{p_1}{\phi_1(\boldsymbol{q}_1)}, \ldots, \frac{p_N}{\phi_N(\boldsymbol{q}_N)}, y\right), i = 1, \ldots, N.$$

The behavioral implications of the pure repackaging framework are intuitive, well known, but sometimes troubling – for example, the individual is indifferent between a doubling of $x_i$ or $\phi_i(\boldsymbol{q}_i)$, and whether an increase in $\phi_i(\boldsymbol{q}_i)$ results in an increase in demand depends critically on whether the price elasticity of demand is less than one in absolute value. From a more practical perspective, an appealing attribute of the pure repackaging approach is that it satisfies the so-called Willig condition (Willig (1978)) which implies that a simultaneous change in $p_i$ and $\phi_i(\boldsymbol{q}_i)$ can be transformed into a pure price change that generates the same level of satisfaction for the consumer. As von Haefen (1999) demonstrates, The Willig condition implies that even when the analyst cannot recover a closed form representation of the full structure of consumer preferences from observable Marshallian demands, she can nevertheless use Vartia's (1983) numerical algorithms to construct exact welfare measures for simultaneous price and quality changes.

A second commonly used empirical specification consistent with (4) and (5) is the cross-product repackaging approach introduced by Willig (1978). Preferences in this case can be nested within the following general structure:

$$(9) \qquad U\left(\boldsymbol{x}, z + \sum_{i}^{N} \delta_i(\boldsymbol{q}_i) x_i\right).$$

where $\delta_i(\boldsymbol{q}_i), i = 1,...N,$ are commonly referred to as cross-product repackaging parameters. Assuming interior solutions for $\boldsymbol{x}$ and $z$, the indirect utility and Marshallian demand functions in this case can be written generally as:

$$(10) \qquad V\left(p_1 - \delta_1(\boldsymbol{q}_1), ..., p_N - \delta_N(\boldsymbol{q}_N), y\right),$$

$$(11) \qquad x_i = g_i\left(p_1 - \delta_1(\boldsymbol{q}_1), ..., p_N - \delta_N(\boldsymbol{q}_N), y\right), \; i = 1, ..., N.$$

As Hanemann (1984) suggests, the cross-product repackaging approach is in many ways less restrictive than the pure repackaging approach – the consumer is not necessarily indifferent between a doubling of $x_i$ or $\phi_i(\boldsymbol{q}_i)$, and so long as the elasticity of demand is strictly negative, Marshallian demand will rise with a quality improvement. Like the pure repackaging approach, the cross-product repackaging approach implies Marshallian demand functions that are consistent with the Willig condition. As a result, exact welfare measures for simultaneous price and quality changes can be constructed using Vartia's numerical techniques regardless of the existence of closed form representations of preferences.

One feature of the cross-product repackaging framework that may limit its empirical usefulness is the possibility of negative quality adjusted prices (i.e., $(p_i - \delta_i(\boldsymbol{q}_i)) < 0$) and its implications for consumer behavior.[18] The nature of the problem can be appreciated by studying the first order conditions implied by the consumer's problem for a simple two good model where preferences can be represented by the utility function $U\left(x, z+\delta(\boldsymbol{q})x\right)$. If utility is strictly increasing in $(x, z)$ and $(p-\delta(\boldsymbol{q}))<0$, then

$$(12) \qquad U_1 - U_2(p-\delta(\boldsymbol{q}))>0,$$

where $U_i$ is the derivative of the direct utility function with respect to its $i$th argument. Equation (12) implies that a negative quality adjusted price results in the consumer spending all of her income on $x$.[19] Although possible, this outcome is extreme and unlikely to be consistent with micro data. Of course, it is only a concern to the degree that negative quality adjusted prices arise in practice which will vary from application to application. The outcome, however, is more likely to arise when prices are relatively small and individual preferences with respect to quality vary substantially. In these cases, the cross-product repackaging framework may be an undesirable approach for developing weakly complementary preferences.

A third approach for developing weakly complementary preferences was suggested indirectly by Larson (1991) and HKP and shall be referred to as the generalized translating approach. A generic direct utility function that encompasses specifications consistent with this approach is:

$$(13) \qquad U\left(\zeta_1(x_1,\boldsymbol{q}_1)-\zeta_1(x_1=0,\boldsymbol{q}_1), \zeta_2(x_2,\boldsymbol{q}_2)-\zeta_2(x_2=0,\boldsymbol{q}_2), \ldots, \zeta_N(x_N,\boldsymbol{q}_N)-\zeta_N(x_N=0,\boldsymbol{q}_N), z\right)$$

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[18] A similar problem can arise when quality adjusted prices are negative with the pure repackaging approach. However, by using transformations that restrict $\phi_i(\boldsymbol{q}_i)$ to be strictly positive, these difficulties can be avoided altogether.

[19] This result carries over to the case where $x$ is a vector and only one quality adjusted price is negative. The solution to the case where more than one quality adjusted price is negative is more complicated to characterize generally, but it will always be the case that $z$ will not be consumed.

where each $\zeta_i(x_i, \boldsymbol{q}_i)$ subfunction is strictly increasing in $x_i$. Loosely speaking, the $\zeta_i(x_i = 0, \boldsymbol{q}_i)$ terms serve as translating parameters (Pollak and Wales (1992)) that jointly translate or shift the $\zeta_i(x_i, \boldsymbol{q}_i)$ terms and in turn the consumer's indifference curves in ways that result in weak complementarity holding. Because of this property, $\zeta_i(x_i = 0, \boldsymbol{q}_i)$ is referred to here as a generalized translating parameter.

In general, little can be said about the indirect utility and Marshallian demand structures implied by (13) without placing additional structure on $\zeta_i(x_i, \boldsymbol{q}_i)$. It is uncertain, for example, whether consumption will be strictly increasing in quality or the Willig condition will be satisfied. This suggests that the analyst should carefully study the behavioral and welfare theoretic properties of an empirical specification to insure their plausibility in a given application.

Combined, the pure repackaging, cross-product repackaging, and generalized translating approaches represent three viable repackaging strategies for developing weakly complementary demand models for applied researchers. In addition equation (4) above suggests that several more general repackaging approaches with potentially more appealing implications for behavior and preferences are available to applied researchers. A concrete example may be instructive. Consider the following direct translog specification,

$$U = \sum_i \alpha_i \ln(x_i + b_i) + \sum_i \sum_j \beta_{ij} \ln(x_i + b_i) \ln(x_j + b_j) + \ln(z + b_z).$$

Although it is straightforward to develop pure repackaging, cross-product repackaging, and generalized translating specifications consistent with this structure, other specifications are possible such as

$$U=\sum_i \alpha_i \ln(x_i+b_i(\boldsymbol{q}_i))-\alpha_i \ln(b_i(\boldsymbol{q}_i))+\sum_i\sum_j \beta_{ij}\ln(\tau_i(\boldsymbol{q}_i)x_i+b_i)\ln(\tau_i(\boldsymbol{q}_i)x_j+b_j)+\ln(z+b_z)$$

and

$$U=\sum_i \alpha_i \ln(\tau_i(\boldsymbol{q}_i)x_i+b_i)+\sum_i\sum_j \beta_{ij}(\ln(x_i+b_i(\boldsymbol{q}_i))-\ln b_i(\boldsymbol{q}_i))(\ln(x_j+b_j(\boldsymbol{q}_j)-\ln b_j(\boldsymbol{q}_j))$$
$$+\ln(z+b_z)$$

Finally, it is worth noting that if the analyst is working within the primal framework, developing weakly complementary empirical specifications is relatively straightforward. Beginning with any direct utility function that permits corner solutions and is nested within the general structure $U(f_{11}(x_1),...,f_{1M_1}(x_1),f_{21}(x_2),...,f_{2M_2}(x_2),...,f_{N1}(x_N),...,f_{NM_N}(x_N),z)$, the analyst should replace the $f_{ij}(x_i)$ functions with $f_{ij}(x_i,\boldsymbol{q}_i)$ that satisfy the property that $\partial f_{ij}(x_i,\boldsymbol{q}_i)/\partial \boldsymbol{q}_i=0$ if $x_i=0,\forall i,j$. Within the dual framework, developing pure and cross-product repackaging approaches is straightforward (see equations (7) and (10) above), but generalized translating and other repackaging specifications are more difficult to develop in general.

*3.2 The Integrating Back Approach*

Larson (1991) introduced an alternative and very general strategy for developing weakly complementary empirical models that builds on an approach suggested by Hausman (1981) and LaFrance and Hanemann (1989). He assumes that the analyst begins with an integrable Marshallian demand system where goods' quality attributes enter arbitrarily. Duality theory implies that the following equalities hold:

$$(14)\qquad \frac{\partial E(\boldsymbol{p},\boldsymbol{Q},\bar{U}^0)}{\partial p_i}=x_i=g_i(\boldsymbol{p},\boldsymbol{Q},y)=g_i(\boldsymbol{p},\boldsymbol{Q},E(\boldsymbol{p},\boldsymbol{Q},\bar{U}^0)),\forall i.$$

In some cases one can use the techniques of differential calculus to solve (14) for the closed form quasi-expenditure function $\tilde{E}(\boldsymbol{p},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U}^0))$, or the expenditure function defined in terms of prices, quality, and a constant of integration $k(\boldsymbol{Q},\bar{U}^0)$ that depends on quality and baseline utility. Because $\tilde{E}(\boldsymbol{p},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U}^0))$ is an incomplete characterization of consumer preferences with respect to quality, it can not be used to evaluate the welfare implications of policies that involve quality changes without the analyst placing additional structure on $k(\boldsymbol{Q},\bar{U})$. Larson recognized, however, that weak complementarity places additional structure on $k(\boldsymbol{Q},\bar{U})$ that may facilitate welfare measurement for quality changes. When $\tilde{E}(\boldsymbol{p},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U}^0))$ is evaluated at the Hicksian choke price for the *i*th good $\hat{p}_i(\boldsymbol{p}^{-i},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U}))$, or the price that drives the Hicksian demands for the *i*th quality differentiated good to zero, weak complementarity implies that, regardless of whether the other goods are consumed, the individual is not willing to pay for marginal improvements in the quality attributes of the *i*th good, i.e.,

$$(15) \qquad \frac{\partial \tilde{E}(\hat{p}_i(\boldsymbol{p}^{-i},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U})),\boldsymbol{p}^{-i},\boldsymbol{Q},k(\boldsymbol{Q},\bar{U}))}{\partial \boldsymbol{q}_i} = 0, \forall i.$$

Equation (15) places restrictions on $k(\boldsymbol{Q},\bar{U})$ that can in principle be used to identify its structure up to a constant of integration that only depends on the baseline utility level, i.e., $\tilde{k}(\bar{U})$. As Hausman (1981) has argued, $\tilde{k}(\bar{U})$ can be interpreted as a monotonic transformation of utility, and thus the analyst can arbitrarily set it equal to $\bar{U}$ (i.e., $\tilde{k}(\bar{U}) = \bar{U}$) with no loss in generality. As a result, the analyst has recovered the full structure of preferences with respect to quality.

Although Larson's integrating back approach is irrefutable in its logic, two factors call into question the usefulness of the approach as a general strategy for developing weakly complementary empirical demand models. In his original paper, Larson used simple two good linear demand and linear expenditure models to illustrate the potential usefulness of the approach. A careful inspection of how quality enters each specification suggests that both can be interpreted as special cases of the repackaging approach.[20] Moreover, there have been no multi-good empirical applications that develop weakly complementary demand models via the integrating back approach since Larson suggested the approach over a decade ago. Consequently, there is little evidence that the integrating back approach offers additional insights into how applied researchers can develop weakly complementary demand models.

One approach to evaluating the empirical usefulness of the integrating back approach is to consider a large number of commonly used empirical demand specifications with quality entering in a variety of ways. If the integrating back approach suggests new weakly complementary demand models that could not have been derived from the repackaging approach, then its value to applied researchers is confirmed. Towards this end, 24 different single equation linear, semi-log, and log-linear specifications with quality allowed to enter in alternative ways are considered. A third of the specifications treat demand (i.e., $x$), another third treat expenditure ($e = px$), and the final third treat expenditure share ($s = px/y$) as the dependent variable.[21] All of these specifications or their logarithmic transformations share a simple linear in parameters structure and have been used or suggested in applied demand analysis. A linear quality index was allowed to enter through the constant, price, or income parameter separately for each model and the mechanics of the integrating back approach was used to determine if closed form solutions for weakly complementary preferences could be recovered.

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[20] Larson's linear demand specification was $x = \alpha + \beta p + \delta q + \gamma y$ which can be rewritten as $x = \alpha + \beta(p + (\delta / \beta)q) + \gamma y$. Comparing this to (10) above suggests that it is consistent with the cross-product repackaging approach to introducing quality. Likewise, the weakly complementary direct utility function implied by Larson's linear expenditure system example is $U(x, \boldsymbol{q}, z) = \Psi(\boldsymbol{q})\ln(x + c) - \Psi(\boldsymbol{q})\ln(c) + \ln(z + b)$ which is consistent with the generalized translating approach.

[21] All of these models assume there is a second Hicksian composite good $z$ that is always strictly consumed.

For brevity the results for all 72 specifications are summarized here and reported in their entirety in a technical appendix available from the author upon request. The key finding was that the integrating back approach could be used to recover closed form weakly complementary preference specifications for 36 of the 72 models in general, although 10 of the 36 weakly complementary specifications could have also been generated by either the pure or cross-product repackaging approaches.[22,23] These findings suggest that the integrating back approach does in fact have genuine value to applied researchers by expanding the menu of weakly complementary single equation models available.

An important caveat should be appended to this statement, however. In most applied situations, the researcher is concerned with developing weakly complementary models for a system of goods, and in these more general cases the marginal value of the integrating back approach is far more dubious. The key difficulty is that the restrictions on the constant of integration implied by weak complementarity will in general depend on the combination of other goods consumed in strictly positive quantities as well as their prices. To the degree that these restrictions depend on consumed goods' prices, restrictions on the constant of integration necessary for weak complementarity to hold will not exist.[24] Moreover, if these restrictions

[22] Of the 26 weakly complementary specifications that could not have been generated by the pure or cross-product repackaging approaches, it is possible that some or all could have been generated by other repackaging approaches. Determining whether this is the case would require one to derive the closed form direct utility functions. Due to time constraints, these tedious derivations were not attempted, but it is unlikely that closed form direct utility functions frequently exist.

[23] Interestingly, 4 of the 36 specifications that could not be linked back to closed form weakly complementary preferences have quality entering in ways that are consistent with either the pure or cross-product repackaging approaches. As discussed in the previous section, Vartia's numerical algorithm can thus be used with these specifications to derive exact welfare measures for price and quality changes.

[24] An example may clarify this point. Consider the demand system:

$$x_1 = \alpha_1(q_1) + \beta_{11}p_1 + \beta_{12}p_2$$
$$x_2 = \alpha_2(q_2) + \beta_{12}p_1 + \beta_{22}p_2$$

depend on the combinations of goods consumed but not their prices, the underlying preference ordering will be discontinuous in $\boldsymbol{x}$. As discussed in the following section, discontinuities in consumer preferences have behavioral implications that significantly call into question the usefulness of the integrating back approach to applied researchers.

In sum, these findings suggest that although some useful weakly complementary single equation specifications may arise from the integrating back approach, it does not represent a generic strategy that can consistently generate new and useful weakly complementary empirical demand system specifications. From a practitioner's perspective, the repackaging approach is far easier to work with and holds greater promise in terms of generating useful empirical specifications.

### *3.3. The Discontinuity Approach*

Bockstael, Hanemann and Strand (1986) suggested a third approach for developing weakly complementary demand models that exploits discontinuities in consumer preferences. Similar to traditional discrete choice models, the discontinuity approach's central building blocks are conditional indirect utility functions which are uniquely defined in terms of which of the $N$ quality differentiated goods are consumed in strictly positive quantities. Since there are $2^N$ possible combinations of goods that are either consumed or not consumed, there are in principle $2^N$ conditional direct utility functions, $U_\omega(\cdot)$, where $\omega$ indexes regimes. Because each $U_\omega(\cdot)$ is by assumption only a function of the prices and quality attributes of the goods consumed in

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where the demand equation for the strictly positive Hicksian composite good is suppressed. Using the integrating back approach in this situation suggests that if $x_1 = 0$ but $x_2 > 0$, the constant of integration must equal $\alpha_1'(q_1)(\alpha_1(q)/\beta_{11} + (\beta_{12}/\beta_{11})p_1)$ for weak complementarity to hold, but this is internally inconsistent because the constant of integration is by assumption independent of prices.

strictly positive quantities (i.e., $\partial U_{\omega} / \partial q_j = 0$ if $x_j = 0$), the preference ordering is consistent with weak complementary. The unconditional direct utility function takes the form:

$$(16) \qquad U(\boldsymbol{x},\boldsymbol{Q},z) = \max_{\omega\in\Omega}\{U_{\omega}(\boldsymbol{x}_{\omega},\boldsymbol{Q}_{\omega},z)\},$$

where $\Omega$ encompasses the full set of $2^N$ regimes, $\boldsymbol{x}_{\omega}$ is a subset of $\boldsymbol{x}$ with each element strictly positive, and $\boldsymbol{Q}_{\omega}$ only includes the quality attributes for the goods included in $\boldsymbol{x}_{\omega}$.

Although intuitive, the discontinuity approach suffers from a fundamental difficulty that casts doubt on its usefulness for applied researchers trying to develop weakly complementary demand models. In the context of a simple two good $(x,z)$ model, Figure 2 illustrates the nature of the problem. The figure is based on the utility function

$$(17) \qquad U(x,\boldsymbol{q},z) = \begin{cases} \Psi(\boldsymbol{q})\ln(x+\theta)+\ln z & \text{if } x>0 \\ \ln z & \text{if } x=0 \end{cases},$$

where $\theta > 1$ and $\Psi(\boldsymbol{q}) > 0$. As in Figure 1, two indifference curves corresponding to the same level of utility but different levels of quality ($\boldsymbol{q}^0 > \boldsymbol{q}^1$) are drawn in $(x,z)$ space. Notice that although both indifference curves intersect the $z$ axis at point A, they do not "fan" from point A as in Figure 1. This feature arises because when the individual moves from consuming none of to a infinitesimal small quantity of $x$ holding $z$ and $\boldsymbol{q}$ constant, she receives a large welfare gain. As compensation for this gain, the individual is willing to forego a significant amount of $z$, which explains why point A is significantly above the points where the two indifference curves approach the $z$ axis. This feature of preferences suggests that it is never rational for the individual to completely forego the consumption of $x$. The individual can always be made better off by consuming at least some infinitesimally small quantity of $x$ than by completely foregoing it. Thus a strictly positive quantity of $x$, however small, becomes an essential component of a utility maximizing bundle, which is at odds with the non-essentiality condition required for weak complementarity to hold.

In principle the difficulties associated with the discontinuity approach can be avoided by placing additional structure on consumer preferences. One possibility involves imposing a minimum consumption level for $x$, say $\underline{x} > 0$, if any of the good is to be consumed at all. In Figure 2, such a minimum consumption level, in combination with a budget constraint corresponding to line 1, would imply that the individual would prefer to consume none of the good at all. Although this resolves the essentiality problem with the discontinuity approach, it implies a more complicated model of consumer choice that in practice would be more difficult to estimate. In particular, the necessary conditions for an individual to rationally choose not to consume $x$ in this context are twofold: 1) $\Psi(\boldsymbol{q})/(\underline{x}+\theta) \leq p/(y-p\underline{x})$ and 2) $\ln y > \Psi(\boldsymbol{q})\ln(\underline{x}+\theta)+\ln(y-px)$. If preferences were continuous (i.e., $U(x,\boldsymbol{q},z) = \Psi(\boldsymbol{q})\ln(x+\theta)+\ln z$), however, there would be only one necessary condition ($\Psi(\boldsymbol{q})/\theta \leq p/y$). From a practitioner's perspective, the added complexity associated with deriving estimable empirical models consistent with conditions 1) and 2) are significantly greater than traditional continuous demand models. Thus the approach, while feasible in theory, is probably less useful in practice.

3.4 *Summary of Alternative Approaches*

The above discussion has several implications for applied research. Perhaps the most significant is that, among the three approaches considered, the repackaging approach offers the most helpful guidance to applied researchers wanting to develop weakly complementary demand models. The approach is also flexible and easy to implement when working within the primal framework. Moreover, the existing repackaging approaches that applied researchers have considered – i.e., the pure repackaging, cross-product repackaging, and the generalized translating approaches - by no means exhaust all of the possible structures that analysts can exploit.

## 4. Empirical Comparison – Alternative Specifications & Estimation Strategy

Having discussed the conceptual advantages of the alternative strategies in the previous section, a more practical question is whether they generate qualitatively different policy inference in an applied setting. As a first step toward answering this question, this section outlines the empirical specifications used to compare alternative approaches to developing weakly complementary specifications in the context of so-called "Kuhn-Tucker" models (Wales and Woodland (1983)), or continuous demand system models specified in the primal framework. Because the direct utility function is the point of departure for Kuhn-Tucker models, weakly complementary specifications derived via the integrating back approach are not considered. In addition, weakly complementary discontinuous specifications, which do not easily admit a closed-form likelihood function conditional on a vector of estimable parameters when minimum consumption thresholds are included, are also not considered. Thus all of the weakly complementary models that are compared fall within the repackaging approach. This implies that the empirical comparison is somewhat limited in scope, but the discussion in the previous section argued that the repackaging approach is by far the easiest to implement, most flexible, and most widely used of the three approaches. Moreover, the comparison encompasses examples of all approaches to developing weakly complementary demand models that have been previously used in multi-good Kuhn-Tucker demand system applications and is valuable to the degree that it informs applied researchers whether existing approaches generate qualitatively different policy inference.

All of the weakly complementary specifications included in the comparison are variations of the linear expenditure system:

$$(18) \quad U(\boldsymbol{x}, z) = \ln z + \sum_i \left[ \Psi_i \ln(\phi_i x_i + \theta_i) + C_i \right],$$

where $[\Psi_i, \phi_i, \theta_i, C_i]$ are functions whose arguments vary across the alternative specifications.[25] The additively separable structure embedded in (18) restrictively implies that all goods are Hicksian substitutes and have non-negative Engel curves. For the purposes of evaluating the empirical implications of the alternative strategies for incorporating weak complementary under consideration, however, the additive separability assumption should not invalidate the comparison.

Seven separate specifications consistent with (18) are considered and summarized in Table 1. The first is the pure repackaging approach, while the second and third are variations of the generalized translating approach with the former employed by HKP. The fourth thru seventh specifications do not embed weak complementarity and are presented mainly for comparison purposes as well as to illustrate the problems arising with decomposition approaches. Although not presented here, an eighth cross-product repackaging specification was also considered. With this specification, however, quality adjusted prices were occasionally found to be negative. As discussed in section 3, such prices imply that the individual wishes to spend all of her income on quality adjusted goods. Due to the implausibility of this prediction, the specification was dropped from the comparison.

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[25] More general additively separable specifications used by von Haefen, Phaneuf, and Parsons (2004) were also considered and found to generate qualitatively similar results.

Because all seven specifications assume that each $\Psi_i$ can be decomposed into two parts, $\Psi_i^*$ and $\varepsilon_i$, where $\Psi_i = \exp(\Psi_i^* + \varepsilon_i)$, the first order conditions that implicitly define the optimal consumption bundle can be rewritten as:

(19) $\varepsilon_i \leq -\Psi_i^* + \ln(p_i / \phi_i) - \ln(y - \boldsymbol{p}^{\mathrm{T}}\boldsymbol{x}) + \ln(\phi_i x_i + \theta_i), \ \forall i$.

Assuming that each $\varepsilon_i$ can be treated as an iid draw from the type I extreme value distribution with common scale parameter $\mu > 0$, the likelihood of observing $\boldsymbol{x}$ conditional on a vector of estimable parameters is

(20) $l(\boldsymbol{x}) = |\boldsymbol{J}| \prod_i \left[ (\exp(-g_i / \mu) / \mu)^{1_{x_i > 0}} \exp(-\exp(-g_i / \mu)) \right]$

where $g_i$ refers to the right hand side of (19) and $\boldsymbol{J}$ is the Jacobian of transformation. As noted by HKP, a notable feature of this likelihood is that the $C_i$ functions do not enter. As a result, specifications 2, 4, and 6 as well as 3, 5, and 7, which differ only in terms of how the $C_i$ functions are structured, are observationally equivalent in terms of estimation. This feature illustrates the point made earlier that weak complementarity is not a testable restriction and implies that only three separate specifications are estimated in this application.

To flexibly account for unobserved heterogeneity in preferences, all structural parameters are assumed to be normally distributed with unrestricted covariance matrix. This specification generalizes previous applications (e.g., von Haefen, Phaneuf, and Parsons, von Haefen (2004)) where for computational tractability only the parameters entering the $\Psi_i$ functions were assumed to vary randomly across the population.[26] In all three estimated models, a total of 23 randomly distributed parameters enter the model, implying that 299 mean and covariance parameters must be estimated. Within a frequentist or classical framework, estimating such a large number of parameters using maximum simulated likelihood techniques (Gourieroux and Monfort (1996)) would represent a formidable if not prohibitively difficult econometric task.

To avoid these computational difficulties, the approach pursued in this paper is to abandon the frequentist paradigm and work within the Bayesian framework (Kim, Allenby, and Rossi (2002)). The conceptual and empirical differences between frequentist and Bayesian approaches are too numerous and subtle to summarize here, but the interested reader can consult Train (2003) for a detailed discussion. It suffices to say that while estimating all three specified models with the data described in the next section within the classical framework was confounded by computational and convergence difficulties, estimation was feasible in a single overnight run within the Bayesian framework for all three models. Moreover, as Train has pointed out, the Bernstein-von Misses theorem implies that the posterior mean Bayesian estimates, interpreted within a classical framework, are asymptotically equivalent to the maximum likelihood estimates assuming the analyst has correctly specified the data generating process. Thus, qualitative statistical inference should be similar whether one is working in a classical or Bayesian framework assuming one has correctly specified the data generating process and uses a sufficiently large data set.

[26] Specifically, allowing only the $\Psi_i$ parameters to vary randomly implies that the Jacobian of transformation is a function of only fixed parameters and thus need only be recomputed once per observation when simulating the likelihood function. With more general specifications, the analyst must calculate the Jacobian of transformation for every simulation and observation, which substantially increases the computational burden.

In the Bayesian estimation framework, the analyst is assumed to have prior beliefs about the values that a set of parameters $\beta$ can take. These beliefs can be formalized into a prior probability distribution, $p(\beta)$. A set of observations, $\boldsymbol{x}$, that is generated by a process that depends on $\beta$, are then observed, and the likelihood of observing $\boldsymbol{x}$ conditional on alternative values of $\beta$, $l(\boldsymbol{x} \mid \beta)$, can be constructed. Given the likelihood, the analyst updates her prior beliefs about $\beta$. By Bayes' rule, her updated beliefs can be summarized by the posterior distribution, $p(\beta \mid \boldsymbol{x})$. Because $p(\beta \mid \boldsymbol{x})$ often does not have a simple structure whose moments can be easily summarized, Bayesian econometricians have developed a number of sophisticated econometric techniques to simulate from $p(\beta \mid \boldsymbol{x})$.

In this paper, a Gibbs sampling routine in combination with an adaptive Metropolis-Hastings algorithm is used to simulate from the posterior distribution of the continuous demand system's structural parameters. The estimation strategy was first developed by Allenby and Lenk (1994) in the context of mixed logit models but is generic to any situation where the conditional likelihood function has a closed form solution as in equation (20). Diffuse priors for all parameters are assumed to limit the impact the prior distributions have on posterior inference. The basic assumptions and steps of the algorithm are sketched in the technical appendix, and the interested reader should consult Train (2003) for a more detailed discussion.

## 5. Data

The data set used in the empirical application comes from the 1997 Mid-Atlantic Beach Survey conducted by researchers at the University of Delaware. The survey's objective was to measure Delaware residents' beach utilization in the Mid-Atlantic region and its interaction with beach management policies. A county-stratified random sample of Delaware residents were questioned about their visits to 62 ocean beaches in New Jersey, Delaware, Maryland, and Virginia during the past year. After data cleaning, a total of 540 completed surveys remained and are the focus of the empirical application here. *PCMiler* was used to calculate round trip travel times and distances from all 540 individuals' resident zip codes to the 62 beaches, and travel cost measures were constructed assuming that travel time could be valued at the wage rate and the out-of-pocket cost of travel was \$0.35 per mile.

Several beach attributes were collected and are used to represent the quality dimension of beaches in the region. Summary statistics for these characteristics as well as demographic and trip taking information are presented in Table 2. For further details on the data used in this analysis, the interested reader should consult Parsons, Massey, and Tomasi (1999) and von Haefen, Phaneuf, and Parsons (2004).

## 6. Estimation Results

Table 3 reports a selected set of posterior mean and variance parameter estimates for the alternative specifications estimated.[27] The estimates suggest a number of qualitative conclusions. A comparison of mean posterior conditional log-likelihood values suggest that the statistical fits of the specifications where site quality enters through the $\phi_i$ and $\theta_i$ functions are virtually indistinguishable and noticeably better than the fit associated with the specification where site quality enters through the $\Psi_i$ functions. Since each model has the same number of parameters, this finding would suggest that specifications 1, 3, 5, and 7 may be more reliable for policy purposes based on purely statistical grounds. In addition, a comparison across specifications of all mean posterior variances suggests that there is considerable heterogeneity

[27] The remainder of the estimates are available from the author upon request.

across the population. In particular, the magnitudes of the posterior mean and variance point estimates for all quality attributes suggests that there is a diversity of opinions with respect to whether the alternative attributes make a site more or less attractive to visit.

Looking more closely at the quality parameters, one notices that the sign of the posterior mean estimates for the specification where quality enters through the $\theta_i$ function are generally opposite from the other specifications. Given how $\theta_i$, $\phi_i$, and $\Psi_i$ enter preferences, however, these opposite signs are consistent with the notion that an increase in one of site *i*'s quality attributes will have the same directional impact on aggregate consumer demand. Moreover, for the weakly complementary specifications and specification 4, the directional impact on aggregate consumer utility will be the same, but the directional impact will differ for the non-weakly complementary specifications 5 and 7. This latter fact will be helpful in explaining the welfare results reported in the next section.

## 7. Welfare Results

### *7.1 Policy Scenario*

The policy scenario used to evaluate the alternative specifications considers the erosion of all eleven developed (i.e., non-park) beaches in the Delaware/Maryland/Virginia area to widths of seventy-five feet or less.[28] Such an outcome might result if current state-sponsored beach nourishment programs were abandoned. For this scenario, the key parameter from the econometric models is the coefficient on the narrow beach dummy variable. Consistently across the alternative specifications, the parameter estimates suggested that *ceteris paribus* a narrow

[28] In 1997 one of the beaches was less than 75 feet in width, and none were more than 200 feet in width. For further details on this scenario, see von Haefen, Phaneuf, and Parsons (2004).

beach (i.e., a beach with width of 75 feet and less) was less frequently visited by Delaware residents on average, although some individuals found narrow beaches more attractive.

*7.2 Welfare Results*

Summary statistics for the posterior distribution of the expected compensation surplus from the alternative specifications are presented in Table 4, and the details on how these estimates were constructed are reported in the technical appendix. One of the most striking results is that the mean estimates in panel A are similar in magnitude across specifications 1, 2, and 3, the weakly complementary models. This finding reflects the fact that all three specificationss predict similar changes in total trips for the policy scenario (-1.9541, -2.0150, and -1.9405, respectively) and rule out nonuse values. Collectively, they suggest that welfare estimates are relatively robust to the alternative weakly complementary repackaging specifications that have been previously considered in valuation studies.

Turning to specifications 4 thru 7 that do not assume weak complementarity holds but are behaviorally equivalent to specifications 2 and 3, three sets of results are reported. Beginning with the total value estimates in panel A, one finds qualitatively different welfare estimates relative to the corresponding weakly complementary specifications. For example, specification 6's mean estimate is more than triple the magnitude of specification 2's mean estimate, and specification 7's estimate is over 290 times the absolute magnitude of specification 3's and the opposite sign. As suggested in the previous section, the positive sign of specification 7's mean estimate (as well as specification 5's) arises because of the model's counterintuitive and highly questionable prediction that beach erosion causes utility to rise while demand falls. The estimates for specifications 4 and 5 are similar qualitatively to specifications 6 and 7's estimates although less extreme. Collectively, the estimates in panel A highlight that in general specifications that assume weak complementary will imply qualitatively different welfare estimates from those that do not.

Panels B and C of Table 4 report welfare estimates for specifications 4 thru 7 that exploit variations of the decomposition approach suggested by HKP. As discussed in section 2, these decomposition approaches attempt to isolate the use component of value by subtracting from the total value estimates the nonuse component. Two important judgments are required with this approach: 1) which non-weakly complementary utility function to use and 2) whether the nonuse component of value is defined when the demands for just the affected or all sites are set to zero.[29] Comparing welfare estimates between specifications 4 and 6 and 5 and 7, respectively, suggest the importance of the former assumption, and two sets of results are presented in Table 4 to suggest the importance of the latter – one where nonuse values are measured when only the affected sites' demands are restricted to zero (panel B) and another where all sites' demands equal zero before and after the quality change (panel C). Collectively, these estimates suggest that in general both sets of judgments can have important implications for policy.

Finally, some simulations and algebra not reported here show that the decomposition estimates reported in panels B and C would have generated identical welfare measures to the weakly complementary specifications in panel A if preferences were quasilinear in addition to being additively separable. Thus the reason why divergences among these estimates are found in this application are due to income effects embedded in the linear expenditure system. More generally, to the degree that preferences are not additively separable and quasilinear, one should

[29] HKP likely appreciated these

expect at least some differences between welfare estimates from weakly complementary models and use component welfare estimates from non-weakly complementary models.

## 8. Conclusion

This paper has conceptually and empirically compared alternative empirical strategies for incorporating weak complementarity into continuous demand system models. Three main implications for future research can be drawn from the paper's findings. First, the repackaging approach offers the most guidance to applied researchers when developing weakly complementary demand models. The approach is relatively easy to implement within the primal framework and offers the researcher considerable flexibility that remains largely untapped at present. Second, the empirical results reported in this paper suggest that among the existing menu of repackaging approaches, qualitatively similar policy implications can be expected if preferences are additively separable, but future research should study their properties in the context of non-additively separable models and more general repackaging approaches before conclusions are drawn with confidence. Third and finally, the decomposition approaches suggested by HKP for specifications that do not assume weak complementarity will in general depend on arbitrary assumptions about how to differentiate use from nonuse values. The empirical results suggested that use-related welfare estimates can be very sensitive to these judgments.

One final point is worth emphasizing in closing. Although this paper has relied upon intuitive and practical reasons to suggest why analysts should develop demand system models that are consistent with weak complementarity when only revealed preference data are available, it remains a strong and arbitrary restriction. As HKP have argued, future research should attempt to combine revealed and stated preference data in ways that allow the quality attributes of goods to enter preferences more flexibly when nonuse values are likely present (Cameron (1992)). By doing so, the analyst would in principle be able to test for weak complementarity and, when rejected, recover nonuse values. Moreover, such studies could suggest to applied researchers working with just revealed preference data when welfare estimates derived from weakly complementary models are likely to capture total value. At present, analysts must base their assessment of the appropriateness of weak complementarity entirely on intuition, but empirical evidence suggesting the conditions under which the restriction is likely to hold can only improve the credibility of valuation estimates.

*Table 1*
***Alternative Specifications***

| Specification | Restrictions |
|---|---|
| *1) Pure Repackaging* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i),\ \phi_i = \exp(\delta^{\mathsf{T}} \boldsymbol{q}_i),\ \theta_i = \exp(\theta^*),\ C_i = 0,\ \forall i$ |
| *2) Generalized Translating #1* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \delta^{\mathsf{T}} \boldsymbol{q}_i + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^*),\ C_i = -\theta^* \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \delta^{\mathsf{T}} \boldsymbol{q}_i + \varepsilon_i),\ \forall i$ |
| *3) Generalized Translating #2* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^* + \delta^{\mathsf{T}} \boldsymbol{q}_i),\ C_i = -(\theta^* + \delta^{\mathsf{T}} \boldsymbol{q}_i) \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i),\ \forall i$ |
| *4) No Weak Complementarity #1* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \delta^{\mathsf{T}} \boldsymbol{q}_i + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^*),\ C_i = 0,\ \forall i$ |
| *5) No Weak Complementarity #2* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^* + \delta^{\mathsf{T}} \boldsymbol{q}_i),\ C_i = 0,\ \forall i$ |
| *6) No Weak Complementarity #3* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \delta^{\mathsf{T}} \boldsymbol{q}_i + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^*),\ C_i = (\exp(\tau^{\mathsf{T}} \boldsymbol{s} + \delta^{\mathsf{T}} \boldsymbol{q}_i + \varepsilon_i) \exp(\theta^*))^2,\ \forall i$ |
| *7) No Weak Complementarity #4* | $\Psi_i = \exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i),\ \phi_i = 1,\ \theta_i = \exp(\theta^* + \delta^{\mathsf{T}} \boldsymbol{q}_i),\ C_i = (\exp(\tau^{\mathsf{T}} \boldsymbol{s} + \varepsilon_i) \exp(\theta^* + \delta^{\mathsf{T}} \boldsymbol{q}_i))^2,\ \forall i$ |

where:

$\boldsymbol{s}$ ~ vector of demographic variables

$\varepsilon_i$ ~ unobserved heterogeneity distributed iid type I extreme value

$[\tau, \delta, \theta^*]$ ~ random parameters

For all specifications to insure $\mu > 0$, $\mu^*$ is estimated where $\mu = \exp(\mu^*)$.

***Table 2***
***Sample Demographics and Beach Quality Characteristics***

| *Variable* | *Description* | *Mean (std. err.)*[1] |
|---|---|---|
| *Sample demographics (540 respondents)* | | |
| Ln(age) | Natural log of respondent age | 3.821 (0.334) |
| Kids under 10 | Respondent has kids under 10 (0/1) | 0.267 |
| Kids 10 to 16 | Respondent has kids between 10 and 16 (0/1) | 0.206 |
| Vacation property in DE | Respondent owns vacation home in DE (0/1) | 0.037 |
| Retired | Respondent is retired (0/1) | 0.248 |
| Student | Respondent is student (0/1) | 0.0481 |
| Income | Household annual income | 49,944 (30,295) |
| Trips | Total visits for day trips to all sites | 9.776 (14.07) |
| Sites visited | Number of beaches visited during 1997 | 2.691 (3.199) |
| *Site quality characteristics* | | |
| Beach length | Length of beach in miles | 0.624 (0.872) |
| Boardwalk | Boardwalk with shops and attractions (0/1) | 0.403 |
| Amusements | Amusement park near beach (0/1) | 0.129 |
| Private/limited access | Access limited (0/1) | 0.258 |
| Park | State or federal park or wildlife refuge (0/1) | 0.097 |
| Wide beach | Beach is more than 200 feet wide (0/1) | 0.258 |
| Narrow beach | Beach is less than 75 feet wide (0/1) | 0.145 |
| Atlantic City | Atlantic city indicator (0/1) | 0.0161 |
| Surfing | Recognized as good surfing location (0/1) | 0.355 |
| High rise | Highly developed beach front (0/1) | 0.242 |
| Park within | Part of the beach is a park area (0/1) | 0.145 |
| Facility | Bathrooms available (0/1) | 0.484 |
| Parking | Public parking available (0/1) | 0.452 |
| New Jersey | New Jersey beach indicator (0/1) | 0.742 |
| Travel cost | Travel Cost = (round trip travel distance) ×($0.35) + (round trip travel time)×(wage rate) | $118.42 (51.67)[2] |

[1] Summary statistics for household variables are means (standard errors) over the 540 individuals. Summary statistics for site variables are means (standard errors) over the 62 sites.
[2] This statistic is the mean (standard error) of each individual's mean round trip travel cost. Each individual in the sample has a unique travel cost associated with visiting each of the 62 sites. Since prices are functions of distance, there is substantial variability in travel costs both across individuals and sites.

***Table 3***
***Some Posterior Parameter Estimates***[1]

| | *Specification 1* | | *Specifications 2 & 4* | | *Specifications 3 & 5* | |
|---|---|---|---|---|---|---|
| *Site i's quality attributes enters through* | $\phi_i$ | | $\Psi_i$ | | $\theta_i$ | |
| *Conditional Log-Likelihood* | -4,599.7 | | -5,207.15 | | -4,603.2 | |
| | (62.153)[2] | | (60.793) | | (66.576) | |
| | *mean* | *variance* | *mean* | *variance* | *mean* | *variance* |
| *Quality parameters* | | | | | | |
| Beach length | 0.0581 | 0.4585 | -0.0103 | 0.4349 | 0.0018 | 0.4802 |
| | (0.0709) | (0.0582) | (0.0707) | (0.0682) | (0.0735) | (0.0675) |
| Boardwalk | -0.0229 | 1.1090 | -0.0309 | 1.0738 | 0.2965 | 0.9728 |
| | (0.1029) | (0.2088) | (0.1768) | (0.2159) | (0.1245) | (0.2560) |
| Amusements | 1.9796 | 1.9766 | 1.9230 | 1.9090 | -2.0466 | 2.0519 |
| | (0.1432) | (0.3464) | (0.1531) | (0.4137) | (0.1397) | (0.4969) |
| Private/limited access | -1.1182 | 2.1227 | -1.1681 | 1.7125 | 1.0523 | 1.7836 |
| | (0.1377) | (0.4655) | (0.1601) | (0.3748) | (0.2194) | (0.4416) |
| Park | 0.1717 | 2.0334 | 0.2291 | 1.9011 | -.0203 | 2.4524 |
| | (0.1426) | (0.4349) | (0.1907) | (0.4791) | (0.1799) | (0.5895) |
| Wide beach | -0.8097 | 1.3359 | -0.8193 | 1.1916 | 0.7449 | 1.2932 |
| | (0.1122) | (0.1991) | (0.1163) | (0.2493) | (0.1130) | (0.2182) |
| Narrow beach | -1.6687 | 1.8250 | -1.4698 | 1.5513 | 1.6029 | 1.6742 |
| | (0.2509) | (0.4184) | (0.2464) | (0.4011) | (0.2313) | (0.4624) |
| Atlantic City | 1.3803 | 2.3706 | 1.8875 | 1.6759 | -1.4984 | 2.8793 |
| | (0.2340) | (0.7300) | (0.2219) | (0.3872) | (0.3202) | (0.6266) |
| Surfing | 0.6470 | 1.3630 | 0.6542 | 1.2675 | -0.6109 | 1.3923 |
| | (0.1023) | (0.2420) | (0.1122) | (0.2419) | (0.1007) | (0.2643) |
| High rise | -0.6415 | 1.9916 | -0.4125 | 1.3305 | 0.6171 | 2.1240 |
| | (0.1399) | (0.4108) | (0.1389) | (0.2525) | (0.1651) | (0.3529) |
| Park within | 1.6657 | 2.6970 | 1.8685 | 2.1328 | -1.6215 | 3.2545 |
| | (0.2286) | (0.5480) | (0.3904) | (0.4677) | (0.2796) | (0.7803) |
| Facility | -0.3740 | 0.7551 | -0.3660 | 0.8591 | 0.4977 | 0.8215 |
| | (0.1205) | (0.1236) | (0.1339) | (0.1870) | (0.1220) | (0.1430) |
| Parking | 0.6060 | 1.1607 | 0.9368 | 1.1495 | -0.9775 | 1.1368 |
| | (0.1475) | (0.2717) | (0.2221) | (0.3022) | (0.1172) | (0.2275) |
| New Jersey | -2.9221 | 2.5998 | -3.6821 | 2.9778 | 3.1803 | 2.9481 |
| | (0.1688) | (0.4670) | (0.3027) | (0.8145) | (0.2964) | (0.6932) |
| *Misc. parameters* | | | | | | |
| $\theta^* = \ln\theta$ | 3.0409 | 2.3932 | 0.5929 | 0.3789 | 3.0316 | 2.3384 |
| | (0.1363) | (0.3712) | (0.0749) | (0.0464) | (0.1365) | (0.3811) |
| $\mu^* = \ln\mu$ | -0.8975 | 0.3133 | -0.4841 | 0.2954 | -0.8764 | 0.3128 |
| | (0.0563) | (0.0394) | (0.0583) | (0.0356) | (0.0706) | (0.0417) |

[1] All estimates generated with 50,000 Gibbs sampling iterations. Simulations from the first 25,000 iteration were discarded as burn-in and every 10th simulation thereafter was used in constructing these estimates.
[2] Standard errors across the 2,500 simulations used to construct the point estimates are reported in parentheses. As Train (2003) notes, these can be interpreted as the asymptotic standard errors estimates within a frequentist perspective.

***Table 4***
***Posterior Expected Hicksian Consumer Surplus Estimates for Lost Beach Width at All Delaware/Maryland/Virginia Developed Beaches***

| ***Panel A*** *– Total Value Estimates* | *Mean*[1] | *Std. err.* | *95% credible set* |
|---|---|---|---|
| *1) Pure Repackaging* | -$94.42 | 18.12 | [-$125.69, -$54.52] |
| *2) Gen. Trans. #1* | -$92.95 | 57.35 | [-$159.65, $54.64] |
| *3) Gen. Trans. #2* | -$95.20 | 17.79 | [-$132.04, -$58.47] |
| *4) No Weak Comp. #1* | -$255.77 | 84.63 | [-$382.31, -$59.59] |
| *5) No Weak Comp. #2* | $2,773.80 | 353.77 | [$2,167, $3,561] |
| *6) No Weak Comp. #3* | -$327.88 | 123.59 | [-$567.87, -$68.38] |
| *7) No Weak Comp. #4* | $27,831.57 | 3,589.4 | [$20,178, $34,780] |
| ***Panel B*** *- Decomposition Approaches – nonuse value component of total value arising when demands at only affected sites are restricted to zero before and after quality change* | | | |
| *4a) No Weak Comp. #1 – Use Value* | -$96.32 | 52.66 | [-$159.77, $33.58] |
| *5a) No Weak Comp. #2 – Use Value* | -$94.87 | 21.80 | [-$134.66, -$48.45] |
| *6a) No Weak Comp. #3 – Use Value* | -$98.21 | 48.40 | [-$159.62, $20.35] |
| *7a) No Weak Comp. #4 – Use Value* | -$237.08 | 88.13 | [-$441.00, -$115.29] |
| ***Panel C*** *- Decomposition Approaches – nonuse value component of total value arising when demands at all sites are restricted to zero before and after quality change* | | | |
| *4b) No Weak Comp. #1 – Use Value* | -$97.67 | 52.81 | [-$162.00, $34.97] |
| *5b) No Weak Comp. #2 – Use Value* | -$166.44 | 29.79 | [-$218.73, -$101.02] |
| *6b) No Weak Comp. #3 – Use Value* | -$139.07 | 70.08 | [-$284.56, -$3.18] |
| *7b) No Weak Comp. #4 – Use Value* | -$609.23 | 223.24 | [-$1,115, -$318] |

[1] Expectations generated with 4 simulations for each of the 2,500 posterior parameter draws. Sampling weights implied by county stratified sampling designed used in all estimates.

*Figure 1*
*Weak Complementarity Graphically (from Smith and Banzhaf (2004))*

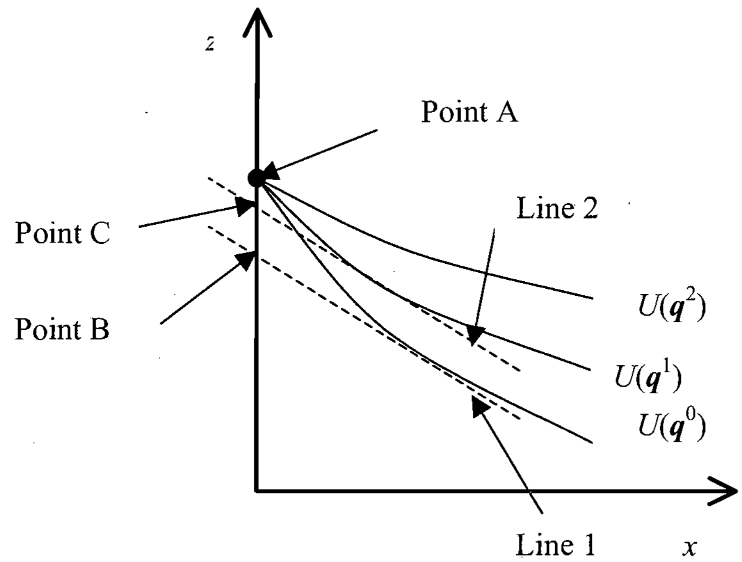
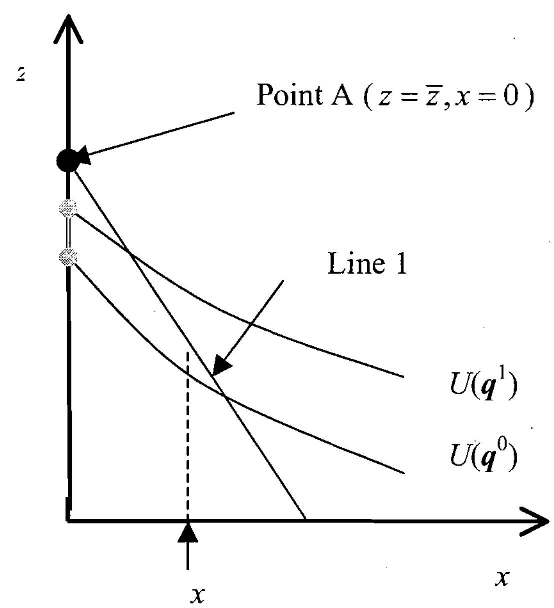



*Figure 2*
*Discontinuity Approach to Imposing Weak Complementarity*



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# Public and Hunter Trade-offs Between Deer Populations and the External Effects of Deer

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## Introduction

### *Deer Population Trends*

White-tailed deer (*Odecoileus virginianus*) populations in much of the U.S. have increased substantially over the last century. Though there is no consensus on the size of the current population in the U.S., estimates have ranged from 15 to 25 million nationwide (McCabe and McCabe 1997). There is, however, consensus among many scientists that in much of their North American range deer population densities currently exceed historical levels that prevailed at the turn of the century (Alverson et al. 1988; deCalesta 1997; Healy 1997; Woolf and Roseberry 1998). Deer population trends in the state of Michigan closely resemble trends at the national level. In 1972 the Deer Range Improvement Program was initiated to improve and acquire deer habitat, with the goal of 1 million deer in the spring herd (Langenau 1994). Since 1972 deer populations have shown a marked increase, and the Michigan Department of Natural Resources (MDNR) currently maintains a white-tailed deer population goal of 1.3 million deer in the fall herd, though the current population is estimated at approximately 1.8 million (Insurance Institute of Michigan).

The deer population increase has created both costs and benefits for Michigan residents. For example, in 2001 there were approximately 753,000 white-tailed deer hunters in Michigan, and slightly over one million people who participated in non-consumptive use activities such as photography and wildlife watching (NSFHWAF, 2001). There may also be an existence value associated with deer, regardless of any consumptive or non-consumptive use. In Michigan, two patent deer-related externalities include deer crop damage and deer-vehicle collisions. Deer-vehicle collisions have increased from 34,352 in 1986 to 67,669 in 1999, with an average insurance claim after a collisions of around $2,000 (Insurance Institute of Michigan) Estimates of deer crop damage range from $13 to $29 per acre, depending on the crop (Campa et al. 1997). Additional deer-related externalities are more difficult to quantify, but may include health related issues, damage to residential property, and damage to commercial and natural forests. Research in northern Michigan has shown that deer browsing can affect the regeneration of Hemlock, Northern White Cedar, and Aspen (Alverson and Waller 1997; Campa et al. 1996; Verme et al. 1986, Frelich et al. 1985) and deer browsing may contribute to changing ecology in northern Michigan's conifer swamps and may change the structure of plant communities in areas of high deer density (Van Deelen 1996).

*Research Objectives*

Previous research has demonstrated that while attitudes toward deer are generally positive (Lauber et al. 2001; Cristoffel and Craven 2000; Diefenbach et al. 1997; Decker and Gavin 1987; Stout et al. 1997; Curtis and Lynch 2001), deer-related externalities including herd health, property damage, crop damage, landscape damage, and deer-vehicle collisions have been cited as concerns (Curtis and Lynch 2001; Decker and Brown 1986; Stout et al. 1993; Sayre et al. 1992; Cristoffel and Craven 2000). Therefore, like most types of resource management, deer management will involve trade-offs. Information that quantifies the types of trade-offs that stakeholders are willing to make among the deer-related attributes will assist managers in setting target populations. In this study we conduct a choice experiment survey to quantify deer management preferences of two stakeholder groups - hunters and the general public. Specifically, the objectives are to:

(a) quantify the relative importance of deer-related attributes to stakeholder groups

(b) quantify the trade-offs different stakeholders are willing to make for changes to the deer population.

The study was conducted in three distinct regions of Michigan: (I) Barry, Eaton, and Calhoun counties in the southwest lower peninsula; (II) Alpena, Oscoda, Montmorency, Alcona, and Presque Isle counties in the northeast lower peninsula; (III) Baraga, Iron, Dickinson, and Marquette counties in the northwest upper peninsula (Figure 1). The three regions differ in a number of ways that are relevant to deer management, such as deer densities, incidence of wildlife disease, deer habitat quality, and agricultural and forestry activities. Socio-economic characteristics also differ among the three regions, with the southwest region having a higher human population density and average per capita incomes than the northeast or northwest.

## Survey Development and Design Challenges

### *Choice Experiment Survey Attributes*

Selection criteria to determine the suite of deer-related attributes for the choice experiment survey included relevance and importance of the attribute to deer management, deer hunters, and the general public. To determine the relevance and importance to deer management, informal interview and discussion sessions were held with deer biologists, ecologists, and other deer management professionals. These initial sessions resulted in a set of deer-related attributes that were then presented to deer hunters and the general public in the form of focus group discussions. The focus groups revealed that while some attributes were straightforward and easily understood, others were open to a variety of interpretation, and thus required more detailed descriptions. Feedback from focus group participants and follow-up discussions with deer management professionals were used to develop attribute descriptions that facilitated quantitative measurement. Descriptions were developed to be relevant to both hunters and the general public, however, hunter surveys included one additional attribute concerning mature bucks (Table 1).

*Informational Needs*

During the survey development process, focus groups and in-person pretest interviews revealed a number of design issues with the potential to hinder respondent's ability to make a meaningful, well-informed choice. For example, focus group discussions revealed that, in addition to the information provided in the attribute descriptions, participants had difficulty making decisions about the attributes without having any baseline knowledge or reference points. Therefore, to help ground respondents and ultimately elicit an informed choice, information about the current status of each attribute was also provided with the descriptions. Regional status quo estimates were readily available only for the annual number of deer-vehicle collisions, obtained from the Michigan Department of Transportation and the State Police. To develop regional status quo estimates for herd health, residential property damage, and deer damage to the forest ecosystem, sources including wildlife biologists, wildlife veterinarians, professional landscape firms, and forest managers from the Division of Forestry at the MDNR, were consulted through informal interviews. Estimates of deer crop damage were developed using a combination of models based on deer-crop damage claims from Wisconsin and published estimates of deer crop damage in Michigan (Wallmo 2003). In describing the status quo, it was emphasized that the information was a regional estimate made by professionals and that the attribute may be higher or lower in certain areas of the region, depending on deer densities and other conditions.

In the survey instrument, each attribute was given approximately one-half to one full page to present the attribute description and status quo information. Following this information, survey respondents were asked about their experience with the attribute and their level of concern associated with any changes to the attribute status quo. These types of Likert scale questions had several purposes. First, asking respondents about individual attribute changes helped prepare them for the choice task questions, where all attribute levels would vary in different deer management scenarios. Second, we expected that the questions would encourage respondents to read each page carefully and discourage page skipping, as well as breaking up the text itself. In addition, these questions allowed us to collect additional preference data on the individual attributes, and on respondent's level of experience and familiarity with deer and deer-related externalities. This type of information can be useful to compare with preference data elicited through the choice experiment and to collect information on the extent of deer-related externalities in the three survey regions.

*Experimental Design*

After developing estimates of the status quo, the initial choice set for the survey instrument consisted of paired comparisons between the status quo and one alternative. However, the pre-test interviews revealed that when comparing alternatives to the status quo, some respondents tended to make choices based solely on deer numbers, without comparing all of the attributes. In an effort to encourage respondents to compare among all the attributes, the choice set was changed to contain three alternatives – an alternative that consisted of the status quo attribute levels, and two alternatives with the same number of deer but varying levels of the other attributes (Figure 2).

A main effects experimental design plan was used to create alternative scenarios to the status quo. However, maintaining independent variation among all the deer-related attributes created some cognitively difficult scenarios for focus group and pre-test participants. For example, relative to the status quo scenario, an alternative scenario may contain more deer but less deer damage to agriculture. Frequently when participants were faced with this type of situation they found the scenarios uncredible and were unable or would not make a choice. Although deer management professionals suggested that there would be factors that allow for these situations, e.g. fencing around agriculture or planting crops that are less desirable deer forage, explaining the mechanisms that might permit these types of 'counterfactual situations' to arise would have required a significant amount of text, in addition to the attribute descriptions and status quo information already provided in the survey. In addition, respondents may have found these mechanisms to be uncredible. To avoid this problem, three types of choice sets were created:

(a) all attribute levels in the alternatives increase relative to the status quo

(b) all attribute levels in the alternatives decrease relative to the status quo

(c) attribute levels can increase or decrease, but changes are marginal relative to the status quo.

In testing the latter type of choice set, we found that focus group and pre-test participants were able to accept scenarios where deer numbers may increase but damage to agriculture may decrease relative to the status quo, as long as the changes were small. Within each type of choice set, a main effects design plan was used, with each attribute taking two levels. In effect, this created three choice experiments and allowed us to maintain independent variation in the design. Figure 3 provides a schematic of the experimental design plan. With three choices per survey, the plan produced sixteen different survey versions per region, for a total of 48 survey versions. The surveys varied by region because the status quo reference levels depended on the region.

## Results and Management Implications

### *Survey Response*

To sample two stakeholder groups, deer hunters and a more general group representative of the Michigan public, the survey sample was drawn from two separate sources, the Michigan Secretary of State (SOS) drivers license database and a database of white-tailed deer hunters maintained by the MDNR. The survey was implemented according to Dillman (2000) guidelines, where contacts included a pre-notification letter, first survey mailing, reminder postcard, second survey mailing, and third survey mailing, unless a reply was received.

The response rate for hunters was 66% (N=1,980). Over 95%f the hunter respondents stated that they hunted. On average, respondents have hunted 26 years, and 16 days during the most recent hunting season. The general public response rate was 62% (N=2,970). We anticipated that the general public sample would contain some hunters, however, the percentage of respondents from the public (SOS) sample who stated that they hunt was approximately 40%, which was considerably high. Although the three study regions were primarily rural areas that may contain more hunters than urban areas, we decided to estimate the general public choice model using only the non-hunters in the SOS sample, allowing us to compare preferences for two distinct groups of hunters and non-hunters.

*Choice Model Results*

Separate choice models are estimated for the licensed hunters from the MDNR sample and for the non-hunters from the SOS sample. Parameter estimates are shown in Table 2. Results show that deer provided positive utility to hunter and non-hunter respondents, and three of the five externalities provided disutility to both groups. Additionally, hunter respondents, whose survey version contained an additional attribute describing mature bucks, derived positive utility from mature buck increases. Externalities that were statistically significant to respondents included reduced herd health, deer-vehicle collisions, and heavy forest browsing. The non-significance of deer damage to agriculture and residential property suggests that these externalities were less important, from the perspective of respondents, than the other three. A status quo dummy variable included in the model was also significant, indicating a significant share of choices for the status quo not otherwise explained by the current deer population and externality levels.

Since Michigan has a diverse and extensive agricultural sector, it is interesting that deer damage to agriculture was not significant in the choice models. This may be a reflection of the fact that less than ten percent of all respondents derived any income from farming activities, and thus deer damage to agriculture may not directly affect the majority of respondents. Conversely, approximately forty percent of respondents had experienced some deer damage to their residential property. However, forty percent of those respondents had changed the types of plants in their yards, suggesting that, although deer damage to residential property may be important, there may be methods for reducing this damage that do not involve reducing the population.

Marginal rates of substitution were calculated for all significant attributes to examine the trade-offs hunters and non-hunters would accept for a 1% increase in the deer and mature buck population, relative to the status quo (Table 3). Results show that hunters will accept greater externality increases for increases in mature bucks than for increases in deer. Hunters also accept greater increases in deer-vehicle collisions and deer browsing than will nonhunters for increases in the number of deer, but the two groups will accept about the same increase in poor herd health for an increase in deer.

To summarize the choice model results, both hunters and non-hunters gain utility from deer, and disutility from three of the five externalities – reduced herd health, deer-vehicle collisions, and heavy deer browsing in the forest. While both groups do gain utility from deer, there are some externality increases which are not offset by more deer, as evidenced by the marginal rates of substitution. In general, hunters will accept greater externality increases to have more deer than will nonhunters, though preferences for herd health are similar for both groups, and they tolerate only small decreases to herd health for a deer population increase. Results demonstrate strong preferences from hunters for mature bucks, and their marginal rates of substitution suggest they will accept almost three times the externality increases for more mature bucks than for deer.

*Management Implications*

The choice experiment results demonstrate the fact that stakeholders can and will make trade-offs among deer-related attributes. This finding is not trivial, as deer management feared at the outset of the project that hunters would consistently prefer more deer regardless of other externalities, and conversely, that the non-hunting public would always prefer fewer damages. The results demonstrate that this is not the case: hunters recognize a point where more deer do not offset the associated externality increases, and non-hunters are willing to accept some level of externalities in order to have a deer population. Results also show that some deer-related externalities are relatively more important than others, for both hunters and non-hunters. For example, herd health, deer-vehicle collisions, and heavy deer browsing in the forest are the least acceptable type of externalities, while neither stakeholder group derived significant disutility from deer damage to agriculture or residential property. In comparing the marginal rates of substitution of hunters and non-hunters, managers may also note that, while hunters will generally accept larger externality increases for deer population increases than will non-hunters, both groups have similar substitution rates for herd health.

For management, the results illustrate several important points. First, our results show that management should consider the preferences of more than just deer hunters when developing policies, as deer provide positive utility to both hunters and nonhunters. Second, the findings suggest that stakeholders may find policy changes and the associated outcomes acceptable if they perceive that they are compensated for any utility losses. For example, for a deer population reduction, respondents may be compensated by an increase in herd health, decreases in deer-vehicle collisions, or decreases in the percent of heavily browsed forest area.

Since policy goals are likely to result in deer population decreases, this type of information may help managers design and implement policies that gain stakeholder approval. Finally, managers may also note that, while hunters will generally accept larger externality increases for deer population increases than will non-hunters, both groups have similar substitution rates for herd health. This suggests that incorporating herd health into policy goals, as well as using the importance of this attribute to design and target outreach efforts, may facilitate acceptance of management strategies. This is especially pertinent considering current bovine TB infections of Michigan deer and growing concerns over the spread of Chronic Wasting Disease (CWD). If management does seek to reduce deer populations, the finding emphasizes the importance of demonstrating to the public the biological relationship between reduced herd health and excessive population size.

The importance of mature bucks relative to general deer numbers among hunter respondents suggests an opportunity to gradually shift hunter demands from quantity (maximum harvestable surplus of bucks) to quality (lower deer numbers, higher buck:doe ratio and more mature bucks). Although this type of shift may not occur easily, as preferences for mature bucks over deer numbers are not universal among hunters, the strong preferences for mature bucks in the three study regions indicate that strategies such as Quality Deer Management may receive support in some areas of the state. One benefit of this demand shift could be improved hunter cooperation in controlling deer numbers through antlerless harvest.

Natural resource management must frequently accommodate competing or conflicting interests. For managers seeking public input to assist with decision-making, choice experiments provide a compatible and realistic tool, as the method places individuals in situations where trade-offs must be made. This research has demonstrated that choice experiments can help inform deer managers in Michigan by providing stakeholder preference structures for deer and the related externalities. Ultimately, this type of information may help managers gain public approval for selected deer management strategies and policy goals.

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Figure 1. Study Areas

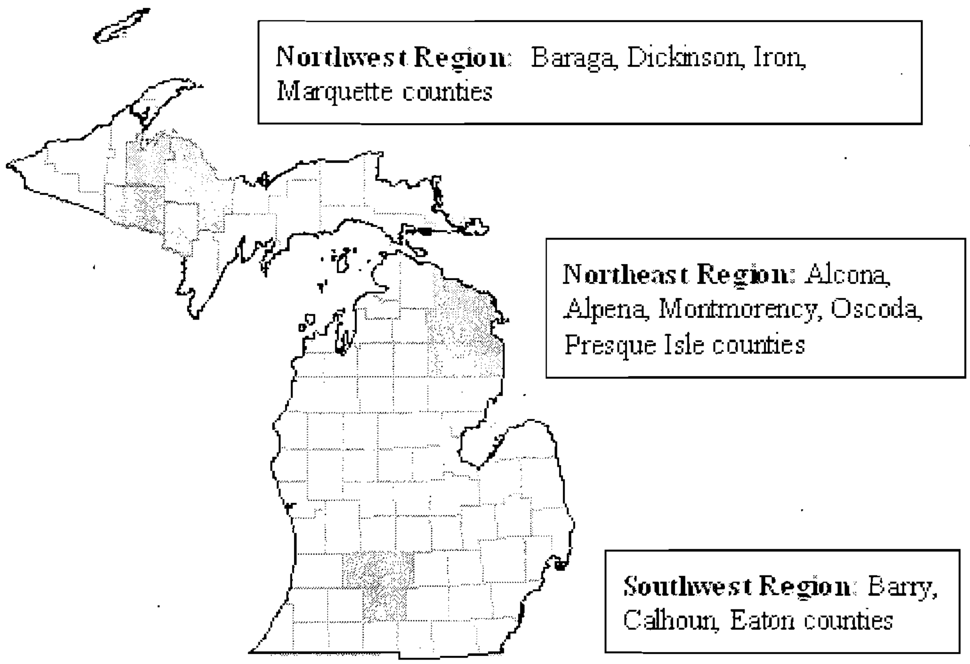

**Figure 2. Choice Set**

| | Current Situation | Scenario A | Scenario B |
|---|---|---|---|
| **Number of Deer** | Current number | 20% more than current number | 20% more than current number |
| **Percent of Deer With At Least One Characteristic of Poor Health** | 25% | 30% | 35% |
| **Percent of Residential Properties Experiencing Deer Damage** | 30% | 35% | 35% |
| **Deer Damage per Acre of Cropland** | $6.00 | $8.00 | $8.00 |
| **Number of Deer-Vehicle Collisions** | 3,562 | 4,600 | 4,300 |
| **Percent of Forest Areas Experiencing Heavy Deer Browsing** | 30% | 40% | 35% |
| Which of these do you prefer? | ❑ Current | ❑ A | ❑ B |

Figure 3. Types of Choice Sets

**A. Increasing Choice Set**

$X_{i}^{\text{StatusQuo}} \;\langle\; X_{i,I}^{\text{Middle Increase}} \;\langle\; X_{i,I}^{\text{Largest Increase}}$

B. Decreasing Choice Set

$X_{i,D}^{\text{Largest Decrease}} \;\langle\; X_{i,D}^{\text{Middle Decrease}} \;\langle\; X_{i}^{\text{Status Quo}}$

C. Marginal Choice Set

$X_{i,M}^{\text{Small Decrease}} \;\langle\; X_{i}^{\text{Status Quo}} \;\langle\; X_{i,M}^{\text{Small Increase}}$

Main effects plan within each choice set to maintain independent variation

$$\mathbf{X}^{\text{largest}} < \mathbf{X}^{\text{middle}} < \mathbf{X}^{\text{smallest decrease}} < \mathbf{X}^{\text{status quo}} < \mathbf{X}^{\text{smallest increase}} < \mathbf{X}^{\text{middle}} < \mathbf{X}^{\text{largest}}$$

**Table 1. Abbreviated Attribute Descriptions**

| Attribute | *Description*[a] |
|---|---|
| Deer | Number of deer in the region |
| Mature Bucks | Number of bucks in region that are two and a half years or older with at least four antler points on one side |
| Herd Health | Percent of deer in region with at least one characteristic of poor health |
| Deer Damage to Residential Property | Percent of residential properties experiencing some deer damage |
| Deer Damage to Agriculture | Deer damage per acre of cropland in region |
| Deer-vehicle Collisions | Annual number of deer-vehicle collisions |
| Deer and the Forest Ecosystem | Percent of forest area in region experiencing heavy deer browsing |

[a] Additional descriptive information was provided in the survey instrument

Table 2. Choice Model Parameter Estimates (std. error)

| Attribute | Hunters | Non-hunters |
|---|---|---|
| Number of deer | 2.6818<br>(0.5667) | 2.1377<br>(0.8037) |
| Number of mature bucks | 7.5800<br>(1.2365) | |
| Reduced herd health | -0.0741<br>(0.0108) | -0.0496<br>(0.0150) |
| Deer damage to residential property | 0.0076<br>(0.0110) | -0.0145<br>(0.0139) |
| Deer damage to agriculture | -0.0537<br>(0.0559) | -0.0583<br>(0.1087) |
| Deer-vehicle collisions | -0.0009<br>(0.0001) | -0.0019<br>(0.0002) |
| Deer browsing in the forest | -0.0097<br>(0.0041) | -0.0310<br>(0.0156) |
| Status quo dummy | 1.0895<br>(0.0475) | 0.8340<br>(0.0521) |
| Log-L | -3361.9 | -1889.7 |

Table 3. Acceptable Trade-offs for a 1% Increase in the Population

| | **Hunters (Mature Buck)** | **Hunters (Deer)** | **Non-Hunters (Deer)** |
|---|---|---|---|
| Percent of deer with at least one characteristic of poor health | 1% | 0.4% | 0.4% |
| Annual number of deer-vehicle collisions | 84 | 30 | 11 |
| Percent of forest experiencing heavy deer browsing | 8% | 3% | 0.7% |