WESTERN REGIONAL RESEARCH PUBLICATION

W-1133
Benefits and Costs of Resource Policies Affecting Public and Private Land

Proceedings from the Annual Meeting
Richmond, VA, March 28-30, 2007

Twentieth Interim Report
August 2007

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Introduction

These proceedings contain selected research papers supporting the 2007 Annual Meeting of the W-1133 Regional Project "Benefits and Costs of Resource Policies Affecting Public and Private Lands", held in Richmond, Virginia, March 28-30, 2007.

The annual convergence of W-1133 scientists representing academia and government took place in Richmond, VA, in 2007. As Richmond celebrated its 400th anniversary, W-1133 celebrated its 40th anniversary with presentations of leading research in valuation and environmental policy and management, an organized panel of federal representatives discussing valuation and policy issues for natural resource-based agencies, and many other exchanges of ideas. The Annual Meetings provide an ideal venue for collaboration in research and sharing of ideas and experiences related to efficient and effective natural resource policy and management. This collection of papers supporting the objectives of W-1133 is illustrative of the breadth and depth of research being applied by the project's members and affiliates.

A panel comprised of federal scientists representing seven agencies was organized to discuss issues and trends in federal agency information needs as they are related to natural resource policy and management. A few recurring themes that arose from the open discussions included the secondary costs of biofuel on agriculture and the environment; climate change; natural disturbances (fire, pests); ecosystem services valuation and provision; and benefit transfer. Our re-chartered project objectives are well-aligned with future information needs of resource policy and management.

I am proud to have been the acting President for the 2006/2007 project-year and to help facilitate some significant milestones. John Loomis and I served as the clearinghouse for research ideas and projects, facilitating the collaborative development of our re-chartering proposal, which has been approved. We also embraced the digital age by converting all past proceedings into electronic form and making this valuable collection of research papers available to everyone on the internet.

The Annual Meeting ran smoothly, and for that I must thank several people. Jerry Fletcher and Roger von Haefen provided AV-support. Klaus Moeltner did so well organizing the social hours last year in San Antonio that I brought him back by popular demand with assistance from Roger. Kelly Giraud (Vice-President) provided invaluable assistance in helping organize the annual meeting. And our thanks is always extended to our advisor, Don Snyder (Utah State University), and our administrative liaison, Fen Hunt (USDA-CSREES), for their ongoing support and guidance in sustaining the W-1133 Project. And a big round of applause and thank you to the many members and friends of W-1133 that make facilitating this group so easy!

Sincerely yours,

Randall S. Rosenberger
Oregon State University

W-1133's Past and Future Project Objectives

2007-2011 (W-2133)

- Natural Resource Management Under Uncertainty
  - Economic Analysis of Agricultural Land, Open Space and Wildland-Urban Interface Issues
  - Economic Analysis of Natural Hazards Issues (Fire, Invasive Species, Natural Events, Climate Change)
- Advances in Valuation Methods
  - Improving Validity and Efficiency in Benefit Transfers
  - Improving Valuation Methods and Technology
- Valuation of Ecosystem Services
  - Valuing Changes in Recreational Access
  - Valuing Changes in Ecosystem Services Flows
  - Valuing Changes in Water Quality

2002-2006 (W-1133)

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

1997-2001 (W-133)

- Valuing Ecosystem Management of Forests and Watersheds;
- Benefits and Costs of Agro-Environmental Policies;
- Valuing Changes in Recreational Access; and
- Benefits Transfer for Groundwater Quality Programs.

1992-1996 (W-133)

- Provide Site-Specific Use and Non-Use Values of Natural Resources for Public Policy Analysis; and
- To Develop Protocols for Transferring Value Estimates to Unstudied Areas.

1987-1991 (W-133)

- To Conceptually Integrate Market and Nonmarket Based Methods for Application to Land and Water Resource Base Services;
- To Develop Theoretically Correct Methodology for Considering Resource Quality in Economic Models and for Assessing the Marginal Value of Competing Resource Base Products; and
- To Apply Market and Nonmarket Based Valuation Methods to Specific Resource Base Outputs.

1974-1986 (W-133)
1967-1972 (WM-59)

Participating Institutions

Auburn University
Colorado State University
Cornell University
Iowa State University
Louisiana State University
Michigan State University
North Carolina State University
North Dakota State University
Ohio State University
Oregon State University
Penn State University
Texas A&M University
University of California Statewide Administration
University of California, Berkeley
University of California, Davis
University of Connecticut-Storrs
University of Delaware
University of Georgia
University of Illinois
University of Kentucky
University of Maine
University of Maryland
University of Massachusetts
University of New Hampshire
University of Rhode Island
University of Wyoming
Utah State University
Washington State University
West Virginia University

List of Attendees, Richmond, VA, 2007

NAME	AFFILIATION
James Caudill	USDOI, Fish and Wildlife Service
Jerry Fletcher	West Virginia University
Kelly Giraud (vice president)	University of New Hampshire
Roger von Haefen	North Carolina State University
Michael Hand	University of New Mexico
LeRoy Hansen	USDA, Economic Research Service
Robert Hearne	North Dakota State University
Daniel Hellerstein	USDA, Economic Research Service
Sandra Hoffmann	Resources for the Future
Tom Holmes	USDA, Forest Service
Wuyang Hu	University of Kentucky
Fen Hunt	USDA, Cooperative State Research, Education and Extension Service
Robert Johansson	US Office of Management and Budget
Robert Johnston	University of Connecticut
Mike Kaplowitz	Michigan State University
Joe Kerkvliet	Oregon State University
Alan Krupnick	Resources for the Future
Andreas Lange	University of Maryland
Linda Langner	USDA, Forest Service
Douglas Lawrence	USDA, Natural Resource Conservation Service
Bob Leeworthy	USDC, NOAA
John Loomis	Colorado State University
Frank Lupi	Michigan State University
Vishakha Maskey	West Virginia University
Dan McCollum	USDA, Forest Service
Don McLeod	University of Wyoming

NAME	AFFILIATION
Klaus Moeltner	University of Nevada, Reno
Ricarda Moser	IASMA
Lara Platt	Cornell University
Greg Poe	Cornell University
Joan Poor	St. Mary’s College
Archana Pradhan	West Virginia University
Alan Randall	The Ohio State University
Richard Ready	Penn State University
Kim Rollins	University of Nevada, Reno
Randall Rosenberger (president)	Oregon State University
Don Snyder (administrative advisor)	Utah State University
Laura Taylor	Georgia State University
Will Wheeler	US Environmental Protection Agency
Lui Xiangph	Grad student

2007 Program

WEDNESDAY, March 28, 2007

5:30p-7:30p	Reception/Social gathering, Manchester Suite, Crowne Plaza Hotel

THURSDAY, March 29, 2007

7:30a-8:00a	**Registration**
8:00a-8:20a	**Presidential Address and Welcome**
8:20a-10:00a	**Session 1: Agro-Environmental Issues** Chair: Robert Johansson (OMB)
8:20-8:40	"Valuing soils information: An assessment of the National Cooperative Soil Survey Program" [Obj #3] *Jerry Fletcher* (WVU)
8:40-9:00	"Willingness to pay for farmland preservation at community and state scales: Implications for benefit transfer" [Obj #4] *Robert Johnston* (UConn), Joshua Duke (UDel)
9:00-9:20	"Farmers' gains from markets for wetland services" [Obj #3] *LeRoy Hansen* (ERS)
9:20-10:00	"Intra-regional amenities, wages, and home prices: The role of forests in the southwest" [Obj #1] *Michael Hand* (UNM), Jennifer Thacher (UNM), Dan McCollum (USFS), Robert Berrens (UNM) Discussant: Don McLeod (UWyo)
10:00a-10:20a	**Break with Refreshments**
10:20a-12:00p	**Session 2: Valuation Methods** Chair: Robert Hearne (NDSU)
10:20-10:50	"A test of proximity as a proxy for environmental exposure in hedonic models" [Obj #3] John Braden (UIll), *Laura Taylor* (GSU), DooHwan Won(UIll)
10:50-11:20	"A new approach to value local recreation using visitors' time allocations" [Obj #2] *John Loomis* (CSU)
11:20-12:00	"Using RP data with alternative specific constants and SP choices to identify preferences in RUM models" [Obj #2] *Roger von Haefen*, Daniel Phaneuf (NCSU) Discussant: Frank Lupi (MSU)
12:00p-1:30p	**Lunch on your own**
1:30p-3:00p	**Expert Panel** Chair: Dan McCollum (USFS) **Issues and Trends in Federal Agency Information Needs** *James Caudill* (USFWS) *Douglas Lawrence* (NRCS) *LeRoy Hansen* (ERS) *Bob Leeworthy* (NOAA) *Fen Hunt* (CSREES) *Will Wheeler* (EPA) *Linda Langner* (USFS)
3:00p-3:20p	**Break with Refreshments**
3:20p-4:40p	**Session 3: Benefit Transfers and Meta-Analysis** Chair: Lara Platt (Cornell)
3:20-3:40	"Forest health values and spatial benefit transfer" [Obj #1] *Tom Holmes* (USFS)
3:40-4:00	"Meta-functional benefit transfer for wetland valuation—Making the most of small samples" [Obj #1] *Klaus Moeltner* (UNR), Richard Woodward (TAMU)
4:00-4:40	"Toward benefit estimates for conservation programs in agriculture—Meta analyses for improvements in wetlands, terrestrial habitat, and surface water quality" [Obj #1] *Alan Randall*, Ayuna Borisova-Kidder, Ding-Rong Chen (OSU) Discussant: Robert Johnston (UConn)
4:50p-5:30p	**Business Meeting—W1133 members only—Manchester Suite**
5:30p-6:30p	**Social gathering, Manchester Suite**

FRIDAY, March 30, 2007

8:00a-10:00a	**Session 4: Stated Preference Valuation** Chair: Randy Rosenberger (OSU)
8:00-8:20	"How much certainty is needed in discrete choice elicitation methods?" [Obj #1] *Lara Platt*, Kent Messer, Gregory Poe (Cornell)
8:20-8:40	"Using bid design and anchoring effect to measure the boundaries of WTP" [Obj #2] *Kim Rollins*, Lucrecia Rodriquez-Barahona (UNR)
8:40-9:00	"Effect of question framing in double bounded contingent valuation of ecosystem services" [Obj #1] *Frank Lupi*, Michael Kaplowitz, Oscar Arreola (MSU)
9:00-9:20	"Correcting for non-response bias in discrete choice models: Using a two stage mixed logit" [Obj #3] *Daniel Hellerstein* (ERS)
9:20-10:00	"Stated preferences for ecotourism alternatives on the Standing Rock Sioux Indian Reservation" [Obj #1] Sheldon Tuscherer, *Robert Hearne* (NDSU) <u>Discussant</u>: John Loomis (CSU)
10:00a-10:30a	**Break with Refreshments**
10:30a-12:00p	**Session 5: Policy Issues** Chair: Vishakha Maskey (WVU)
10:30-11:00	"Regulatory takings, lobby formation and preemption—In the case of Endangered Species Act" [Obj #3] Xiangping Liu, *Andreas Lange* (UMD)
11:00-11:30	"Are biodiversity protections a boon or a bane for local economies? Evidence from the Northwest Forest Plan" [Obj #3] *Joe Kerkvliet* (OSU), Andrew Plantinga (OSU), Henry Eichman (BLM), Gary Hunt (UME)
11:30-12:00	"Testing for intrahousehold bargaining in a stated preference context: Parent's WTP to reduce lead paint hazards for their children" [Obj #3] *Sandra Hoffmann* (RFF), Alan Krupnick (RFF), Vic Adamowicz (UAlberta)
12:00p	**Meeting Adjourns**

W1133 Objectives (2002-2007)
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

PART 1: Papers Supporting Objective:
“Estimate the Economic Benefits of Ecosystem Management of Forests and Watersheds”

On Why Environmental and Resource Economists Should Care about Non-Expected Utility Models

[Forthcoming, *Resource and Energy Economics*

June, 2007]

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Keywords: Environmental Economics, Expected Utility, Risk, Uncertainty

* Senior authorship is shared equally. Shaw holds a joint appointment with Texas A&M's Department of Recreation, Parks and Tourism Sciences and acknowledges their support of this research. In general, we have greatly benefited from having discussions on risk and uncertainty with David Bessler, Trudy Cameron, J.R. DeShazo, Paan Jindapon, William Neilson, Mary Riddel, Antony Scott, and Kerry Smith. We also appreciate helpful comments from the editor of, and two reviewers for, this journal. A considerably earlier version of the paper was presented by Woodward as a selected paper at the 1998 meeting of the American Agricultural Economics Association, August 2-5, 1998, Salt Lake City, Utah. Some of this work was supported by an exploratory grant from the National Science Foundation.

Why Environmental and Resource Economists Should Care about Non-Expected Utility Models

Abstract: Experimental and theoretical analysis has shown that the conventional expected utility (EU) and subjective expected utility (SEU) models, which are linear in probabilities, have serious limitations in certain situations. We argue here that these limitations are often highly relevant to the work that environmental and natural resource economists do. We discuss some of the experimental evidence and alternatives to the SEU. We consider the theory used, the problems studied, and the methods employed by resource economists. Finally, we highlight some recent work that has begun to use some of the alternatives to the EU and SEU frameworks and discuss areas where much future work is needed.

Introduction

The conventional expected utility (EU) model of von Neumann and Morgenstern (1944) is the central framework for the economist's analysis of choice under risk and uncertainty. For example, Graham adopted the EU framework for his classic derivation of a welfare measure under uncertainty, the option price (Graham 1981; see discussion in Ready 1995). Typical environmental and resource economists interested in welfare changes, non-market valuation, or Pareto Optimality, might know little more than this about risk and uncertainty. However, economists have had a good deal to say about modeling risky behavior in recent years and much of this work calls into question the basic framework of decision making under risk that is a pillar of applied analysis. The question we consider here is whether environmental and resource economists would benefit from more attention to these developments. In this manuscript, the larger *we* (hereafter in italics) signifies this target audience of environmental, energy, and natural resource economists. We seek to bridge the gap between general risk/uncertainty economists and those of us most interested in the risk issues that are prominent today among resource economists (Chichilnisky and Heal 1993). As used here "resource economics" is meant to encompass all of the economic issues that relate to environmental policy, natural resource extraction, and energy economics. Issues of interest to resource economists can overlap greatly with those in health economics, agricultural production and welfare, and transportation safety and efficiency.

Almost all decisions that people make involve some level of risk or uncertainty, and uncertainty is particularly pervasive in the area of resource economics. The EU framework has guided most analysis of decisions in situations of risks with known probabilities and the subjective expected utility model (SEU) of Savage (1954) can handle situations where probabilities are not known by the decision maker. These models have been successful not only because of their compelling axiomatic foundations and ability to describe economic choices, but also for the purely practical reason that their mathematical structure facilitates both theoretical and empirical analysis. Whether the probabilities are objective or subjective, the EU and SEU models can be represented as functions that are linear in the probabilities. Some more general or "non" EU (see Machina 1987) frameworks relax the EU assumptions and lead to models that allow the decision maker to place nonlinear weights on the probabilities (e.g., Machina, 1982; 1987 and 1994 or Quiggin, 1993). If we state below that a study finds behavior that is inconsistent the EU theory, we are limiting our discussion to the standard model, not to these more generalized models.

Despite obvious strength in their ties to the theory of choice, the EU and SEU models have some well-known weaknesses. Most critically, using mostly experimental evidence, researchers have often found that individuals often do *not* behave in a manner consistent with the EU models. We focus here on two prominent problems, frequently referred to as the Allais and Ellsberg Paradoxes. The Allais paradox refers to the tendency of individuals to overweight high-consequence low-probability events. The Ellsberg paradox is that people regularly demonstrate an aversion to ambiguity, preferring situations with clear probabilities to those in which the probabilities are uncertain. Since the initial work of Allais (1953) and Ellsberg (1961), many experiments have been constructed to test for Allais and Ellsberg behaviors, and have frequently found it.

Because of such results it is important for us to ask whether – or at least in what circumstances of interest to resource economists – the basic EU/SEU paradigms should be retained? Below we compile a list of a number of issues frequently studied by resource economists. For some such problems (e.g., financial risks in resource management) the primary risks are probably not at the extremes and tend to be well understood. But for many other issues, such as natural catastrophes in the advent of climate change, nuclear accidents, or exposure to toxics, the salient concerns are frequently low-probability high-consequence events about which the uncertainty is great. For example, in the work on nuclear/radioactive waste risks, the experts deem mortality risks to be on the order of 2 in 10 million, while a sample of subjects thought these to be thousands of times higher (see Riddel and Shaw 2006). In other problems, such as the valuation of an endangered species, ambiguity is pervasive. We believe, therefore, that for many of the problems in which *we* are interested, the conditions are exactly those in which the EU and SEU are frequently violated.

Lest the reader be thinking that all is hopeless, there are several generalizations or alternatives to the EU and SEU models, yielding theoretical results consistent with observed behavior. Some have strong normative foundations and can have good predictive power; but no one model fits all situations and behaviors. There are also behaviors that could be explained via relaxation of assumptions that economists usually make, but have little to do with the EU itself. We discuss such issues toward the end of the manuscript, but mainly confine ourselves to ideas related to EU axioms of rationality.

The remainder of the paper is organized as follows. First, the EU and SEU models are briefly presented along with a discussion of some evidence of systematic violations of the models' predictions. We then discuss the implications of these violations for the theory *we* use, the issues *we* study, and the methods *we* employ. In the final section we state some caveats, point to other responses to these issues, and remind readers of situations where we think the EU does work well to predict behaviors.

EU and SEU and the paradoxes of Allais and Ellsberg

To begin, we differentiate between the EU and the SEU, noting that the latter recognizes the importance of subjective or perceived risks. While *we* might typically rely on scientific or experts' assessments of risk, psychologists (e.g. Slovic 1987) usually assume that decisions are based on perceived or subjective assessments of risk rather than "objective ones." Savage (1954) showed that "rational" choices can always be described as arising from an expected utility calculation with the decision maker using his or her personal subjective probabilities. Subjective or perceived risks can be in accord with risk assessments calculated by the scientific community, but they can also be at opposite ends of the spectrum. Naturally, the scientific community may not be in agreement in the first place. When the subjective risks perceived by the public differ from those of the experts, conflict can arise. Table 1 highlights the conflicts that are created in the latter situation, assuming that the policy maker (government) sides with the experts' assessments of risk. We return to these conflicts below.

In both the EU and SEU models, choices are made so as to maximize the expected value of an individual's utility, say,

$$V(\{\mathbf{x},\mathbf{p}\}) = \sum_{i=1}^{n} U(x_i)p_i \qquad (1)$$

where $V(\cdot)$ is the individual's ex ante welfare, $U(\cdot)$ is the individual's ex post utility function, $\mathbf{x}_i$ is a vector of goods and services in the i^{th} state of the world, and p_i is the probability that the i^{th} state of the world will actually occur. The EU formulation of von Neumann and Morgenstern (1944) assumes that the probability vector **p** is known by the decision maker. The SEU formulation of Savage (1954) makes no assumptions about the knowledge of **p**, but instead shows that if a decision maker abides by a set of axioms, she will behave *as if* she has a set of personal or subjective probabilities. Although it does not follow from the axioms of either Savage or von Neumann and Morgenstern, agents are typically assumed to be risk averse in theoretical models, meaning that the *U* is concave in the variable involving risk, which is most often income because the main thrust of research in the past 50 years has focused on financial risks that arise for a variety of reasons. However, in empirical applications (of which there are not that many) the assumption of risk aversion is often ignored for convenience and individuals in some estimating sample are assumed to be risk neutral.

As seen in equation (1), the EU and SEU are linear in the probabilities that characterize risks. In effect, the functions assume that individuals do not have preferences over probabilities themselves, only over the outcomes that result. Put another way, utility is outside, not inside the utility function. A limitation of this structure is explained by Yaari (1987):

> In expected utility theory, the agent's attitude towards risk and the agent's attitude towards wealth are forever bonded together. [However,] at the level of fundamental

> principles, risk aversion and diminishing marginal utility of wealth, which are synonymous under expected utility theory, are horses of different colors. (p. 95).

As we will discuss below, relaxing the assumption that preferences are linear in probabilities can go a long way toward resolving many of the observed violations of the EU framework, though as we will also see below, this alone does not solve all the problems that have arisen in analyzing risk preferences (Machina, 1987).

Despite weaknesses, there are many situations in which the EU models perform well and can be used without reservations. [1] These models are especially credible when probabilities are very well understood and clear to the decision maker. One thinks of a simple two-outcome toss of a fair coin, for example. Similarly, as noted by Loomes (2006b), when one option "transparently" dominates the other, decision makers rarely have difficulties making the "right" choice. A key feature of these models that relates to economic valuation is that an individual should not view or perceive of payoffs for a gamble differently from the risks of that gamble (see Loomes, 2006b). This means that there are many situations when the EU models are perfectly appropriate, or at least where connections to issues that are consistent with the traditional EU are plausible.

In Table 2 we provide a list of issues that we could think of which are important to resource economists. The columns on the right indicate our own assessments of the likely magnitude of the probability being quite small (one documented problem area) and the likelihood that ambiguity is important. As seen in Table 2, there are a number of issues that *we* study where risks are well-characterized; they involve neither extreme probabilities nor significant ambiguity. For these problems the EU models are probably quite adequate frameworks for our analysis. A small risk/probability issue does not in itself mean that the EU will fail. However, if people, no matter how hard we try to communicate these small risks to them, simply cannot process the information in such a way as to know how to make decisions, then the first step toward failure of the EU and SEU has been taken.

Violations of the basic axioms of the EU and SEU models have been identified by many researchers including Kahneman and Tversky (1979), Loomes and Sugden (1982), and others.[2] A classic outcome is preference reversal when encountering similarity, known as the common ratio effect (see Starmer 2000; Loomes 2006b). This literature is too vast to present many cases here, but Starmer (2000) provides a good summary up to the date his paper was published. Our focus is mainly on two problems: the Allais and Ellsberg paradoxes. These two problems get particular attention because they are where the largest body of alternative risk research has been conducted and where the possibilities for improving analysis in resource economics seem to be the greatest.

In his path-breaking analysis, Allais (1953) showed that the assumption that utility is linear in the probabilities is not always satisfied. The Allais paradox is that decision makers often do not treat probabilistic outcomes or events in the standard additive manner as required in (1). Instead, dis-proportionally more weight is given to low-probability high-consequence events. Intuitively, we might easily come up with reasons why the magnitude of the probability would matter when we are making a choice. Such reasons give way to the concepts of a "certainty" effect (i.e., I feel that 99% is virtually certain, so why bother with a policy that pushes the probability from this to 100%? Conversely, I want to be absolutely certain and 99% leaves some doubt, so I care much more about this increase of 1% than I do for the same 1% increase from 50% to 51%.). Allais paved the way for work on the impact of scaling down probabilities associated with lotteries, with the frequent result that such scaling flips preferences, when EU theory says that it should not. Such preferences are inconsistent with the EU because they violate

Savage's independence axiom. Closely related is the possibility that people behave as if they weight probabilities.

Probability Weighting Functions: Some Evidence

There is evidence that people make decisions under conditions of risk as if they are actually weighting probabilities (p) that characterize risks. If a weight, w(p), exists over a spectrum of probabilities, it is possible to map out a probability weighting function (PWF). We more formally introduce these below, but note in this section that when the PWF is non-linear, departures from the original von-Neumann-Morgenstern utility function, (e.g. Quizon, Binswanger and Machina 1984), and alternatives are worth considering. This kind of evidence of non-linear PWFs is consistent with Allais' (1953) concerns, and both the common consequence, and common ratio effects, which are damaging for the basic EU.

To our knowledge, the evidence on the existence of PWFs comes from laboratory experiments which makes sense because uncovering the shape of a PWF for an individual can be somewhat difficult. It can be done using a variety of approaches, with the recent literature favoring the trade-off approach (Wakker and Deneffe 1996)[3]. Empirical work has elicited PWFs for mostly monetary, such as Gonzalez and Wu (1999), and much less frequently life-duration, gambles, with a mixture of outcomes (see Bleichrodt and Pinto 2000; or Wakker and Deneffe 1996. For example, the PWF in the monetary-gamble domain is concave for low probabilities and convex for high probabilities with the weight equivalent to the actual probability ranging between 0.3 and 0.5 (Prelec 1999). Less is known about non-monetary gambles and PWFs.

Wakker and Deneffe (1996) use a trade-off method mostly applied to financial gambles, but part of their analysis evaluates the way that a pool of psychology and PhD students viewed risks associated with surgical operations to cure them of unknown health risks associated with a disease they were to imagine they had. Subjects are not told the exact risks of carrying an unfavorable disease, only that they are young and that their probability of carrying the disease is expected to be smaller than for older people, 66 percent of whom carry the disease. Based on their results, they conclude that a higher risk aversion for life duration is due to stronger deviations from expected utility, rather than due to curvature of the utility function.[4]

Other experiments are set up to see if risk preferences stay constant in a manner consistent with linear probability utility functions. They often do not. For example, Bosch-Domènech and Silvestre (1999) find that risk attitudes vary with the level of the endowment for some of their experimental subjects, revealed by subjects' decisions to insure their incomes. A large number of experimental economics studies evaluate when and how the EU models fail and when they succeed (e.g. Hogarth and Kunreuther 1985). Such experimental efforts are by far the most common empirical approach in the literature that has been previously applied to test for violations.[5] Thorough reviews of this literature are available elsewhere.[6] Experiments have mainly involved choices of financial lotteries, often with incentives given for making a choice consistent with our own understanding of decision making under risks.

Experiments that follow from the work of Allais have found that the EU model tends to predict behavior quite well when the decision problem involves probabilities in the "mid-range." However, when faced with small probabilities, nonlinear weighting "is empirically important in explaining choice behavior" (Harless and Camerer 1994, p. 1285). The basic EU model is also not very good at explaining choices when individuals are faced with losses as well as gains (Quiggin 1993 p. 140), as it does not differentiate between these.

Ambiguity

Ellsberg's paradox relates to uncertainty that cannot be reduced to simple probabilities. This kind of uncertainty, which we well refer to as *ambiguity*, is closely related to Knightian uncertainty (Knight 1921) and Shackle's similar concepts (1952, 1969). Ambiguity involves more than just a PWF. There is no room for true *ambiguity* in the EU models. Implicit in the SEU is a Bayesian approach in which uncertainty over the probability distribution, **p**, can be expressed as second-order probability distributions, which can then be additively reduced to a final probability distribution. However, as Ellsberg (1961) showed experimentally, individuals do not follow this reductionist procedure; ambiguity does affect preferences. Camerer and Weber (1992) describe the nature of the ambiguity problem as follows: Suppose you flip a coin 1000 times and get 500 heads. Then take a second coin and flip it twice yielding one head and one tail. There is no objective reason to believe that the probability of heads using the second coin is any different from that of the first. Yet,

> [m]any people … prefer to bet on the first coin, because they are more confident or certain that the first coin is fair. Ambiguity about probability creates a kind of risk in betting on the second coin – the risk of having the wrong belief. SEU effectively requires that decision makers be indifferent towards such a risk. (p. 326)

This systematic aversion to ambiguity cannot be reconciled with the EU or SEU models. First shown experimentally by Ellsberg (1961), there are many "Ellsberg-type" studies showing the existence of ambiguity aversion in preferences.

Some psychologists today have offered two competing hypotheses for the formation of ambiguity: comparative ignorance (Fox and Tversky 1995; Fox and Weber 2002; Chow and Sarin 2001) and the competence hypothesis (Heath and Tversky 1991; Keppe and Weber 1995). In the first hypothesis, an individual's sense of ambiguity grows when he or she is presented with the fact that there is some set of information that is better than what they have when it is vague:

their vague information conflicts with information that appears to be clear. When experts about risk disagree, individuals can place disproportionate weights on high risk outcomes; Viscusi and his colleagues demonstrate this using health risks posed by cigarette smoke (Viscusi, Magat, and Huber 1999). In contrast, the competence hypothesis allows individuals who are particularly knowledgeable or well informed about something (the outcome of a football game) to overcome any feelings of ambiguity. In the studies by Heath and Tversky (1991) and Fox and Tversky (1995), experiments are used, and controlled bets in two settings reveal tendencies to bet on clear rather than vague probabilities. When individuals exhibit a good degree of ambiguity aversion, they tend to pay to avoid it.

Experimental work on ambiguity has focused on two specific and important issues: the distinction between ambiguity and risk, and the degree of aversion to ambiguity. To distinguish these, we use the term *event risk* to refer to the usual uncertainty about an event or outcome (e.g. the risk of getting cancer from a substance in drinking water) and the term *ambiguity* to refer to uncertainty about the risk itself (e.g., the uncertainty about what the actual probability of getting cancer is). In reviewing the literature on ambiguity, Camerer and Weber (1992) found no correlation between risk attitudes and ambiguity attitudes. They also found the existence of "substantial premiums to avoid ambiguity – around 10% to 20% of expected value or expected probability" (p. 340). Ambiguity, therefore, can not only be separate from event risk, but is economically important.

Finally, in one early, non-experimental, and seminal empirical study Kunreuther (1976) finds such evidence among thousands of people interviewed in earthquake and flood-prone areas, including a pool of both insured, and uninsured respondents. He concludes that neither group's behavior can be explained via the "standard" [EU] framework. While part of the sample does not have enough information, and blame rests there, he concludes that even those who do "frequently behave in a manner inconsistent with what would be predicted by [EU] theory." [Kunreuther, 1976, p. 255]. There are very few such empirical studies.

Some Recent Non-Experimental applications of, and evidence, from Non-EU models

Most of the impetus for non-EU models has come from experimental work because laboratory settings allow sufficient controls to test the EU axioms. It is much more difficult to find empirical evidence (not just theoretical) that supports non-EU frameworks, involving "real-world" phenomena (from natural experiments or survey data on intended, or actual behaviors). Actually proving a violation of the SEU in the real world is much more difficult than in the laboratory; naturally, carefully done empirical work only fails to disprove something.

Nuclear waste storage is a setting where work has been done to explore whether individuals' preferences under risk are consistent with standard EU models. Nuclear waste storage is a long-term proposition: high level radioactive wastes must be stored safely for about 10,000 years before their toxic effects diminish to a safe level through a natural decay process. Using assessments of most physical scientists who have worked on the Yucca Mountain Repository project, the long- and short-term mortality risks appear to be extremely low – on the order of 1 or 2 in 10 million. However, even when told this, subjects in a sample believed such risks to be thousands of times higher than the government scientists, classic evidence of failure to comprehend the extremely low probabilities. In one of the only studies using field survey data that we know of, estimated models that depart from the EU are statistically preferred to ones that are more restrictive. The statistical test involves significance of terms that result when incorporating non-linear probabilities. The preferences of sample individuals in these studies exhibit evidence of ambiguity aversion (Riddel and Shaw 2003; Riddel et al. 2003; Riddel and Shaw 2006).

Another empirical study that provides evidence supporting use of a non-EU framework is Cameron's (2005) study on climate change. Again using survey data, and specific functional forms for indirect expected utility functions that allow for non-linear probability weighting. Cameron finds that individual option prices are affected by the dispersion of an individual's subjective probabilities over future environmental quality. This dispersion term is incorporated using the squared deviation between realized temperature and its ex ante expected value. Such preferences do not arise in the linear-in probabilities EU.

Finally, in Baker et al. (2007), we interviewed a small sample of people who were displaced by the major hurricanes, Katrina and Rita. Subjects were asked to state their subjective risk of a hurricane of similar magnitude to Katrina striking the area they came from (most are from New Orleans), after being told that the experts believe this probability to be around 10%. Even after the subjects were walked through educational material (presented to them on a laptop computer with pictures and easy to read text), the average subjective risk for the sample is close to 50%. Of particular interest in our results is the "unsure" explanatory variable. Respondents were asked after their final report of subjective risk whether they were sure or unsure of their risk estimate. On average, stated risks by individuals who are unsure were 20% higher than those who were more confident. This result is consistent with an aversion to ambiguity as operationalized by Gilboa-Schmeidler's Maxmin Expected Utility which we discuss below.

There are also several theoretical concerns regarding the EU framework, though we are not offering these as "evidence" that the EU fails. For example, consider the problem of insurance choices in the EU framework. Quiggin (1993 pp. 80-81) shows that it would never be

EU-optimal to purchase full insurance if rates are actuarially unfair and the utility function is smooth. Intuitively, this is because, in the neighborhood of certainty, the EU preferences are risk neutral so that at the limit agents would prefer to bear some portion of the risk. While alternative explanations for observed full-insurance in the market are possible, Segal and Spivak (1990) show that nonlinear PWFs preferences can explain the purchase of complete insurance. Quiggin (1993) summarizes the result: "even if first order risk aversion is too weak to induce the purchase of insurance, second order risk aversion [nonlinear weighting of probabilities] may lead to a preference for insurance" (p. 81). The ability to discriminate between preferences over the outcomes and preferences over the probabilities in the problem itself, therefore, can prove useful in understanding observed behavior.

We see, therefore, that in both the laboratory and in a few select survey settings the EU models can not describe *all* behavior of economic agents. Where economists go from there, however, is not transparent. There have been two principal responses to the experimental evidence. The first is to treat the results as "phenomena" instead of "paradoxes" (Howard 1992) which require no more of a re-evaluation of the EU models than people's confusion by optical illusions require us to invent new notions of distance (Varian 1992, p. 194; McFadden, 1999). The second response has been to look for theoretical alternatives to the EU and SEU models that can explain the behavioral patterns found in the experimental work, yet retain much of the theoretical appeal and empirical tractability of the EU and SEU models. We examine two of the alternatives in the next section.

Solutions? Alternatives to the EU models

One might first simply conclude from the above discussion about risks that the EU and SEU will always perform well if researchers just keep pushing people hard enough in risk communication exercises. However, there are several reasons this is not the case. First, as indicated above, scientists may not know risks much better than anyone else, leaving doubt about what they are. An example of this is what is known about the human health risks of ingesting arsenic for lower arsenic concentrations, below 50 parts per billion (see Shaw et al. 2006). In surveys and experiments the research is supposed to be truthful with subjects; analysts are obligated to pass on this scientific doubt. Second, understanding risks may take time. For example, consider the assumption that we make that the following two lotteries (denoted with payoff, probability) are equivalent: (30, ½; 40, ½) and (30, ¼; 30, ¼; 40, ½), which is key to development of defining preference changes (Machina, 2001). Well educated people immediately see the equivalence, but it may take time to convey this to those less well educated. There is a limit to what people can take in and process, and it is well known that respondents to surveys and experimental subjects become fatigued after time, so pushing harder in risk communication may lead to some gain on one hand, but losses in mental acuteness later on, perhaps when payoffs are being evaluated. If people just do not "get" probabilities, they may behave in manners that indicate this.[7]

In the next section we present two alternative risk models that seem particularly promising for use in environmental economics: the rank dependent expected utility model, and the maxmin EU model. Both retain much of the structure of the EU and SEU frameworks, yet are able to explain more of the experimental evidence than these models would.[8] This treatment is by no means exhaustive; we are excluding for example Kahneman and Tversky's (1979) prospect theory, Loomes and Sugden's (1982) regret theory, Viscusi's prospective reference theory (1989), and Loomes' (2006b) perceived relative advantage model (PRAM).

The RDEU model

The rank dependent expected utility model (RDEU) that was originally proposed by Quiggin (1982), has also been explored by Yaari (1987), and is closely related to the models of Schmeidler (1989) and Gilboa (1987). Originally developed to explain the Allais paradox, the RDEU model allows probabilities to enter nonlinearly into an individual's objective function. Because it preserves first-order stochastic dominance, Machina (1994, p. 1237) called the RDEU the first "successful" model with nonlinear probabilities.

The RDEU model (presented in depth in Quiggin 1993), is typically applied in problems with well-defined probabilities, but is not restricted to that context.[9] Following Quiggin (1993), the RDEU model is most easily presented in terms of a single lottery in which there are n possible outcomes, x_i; i=1,…,n, which are ordered from worst (i=1) to best (i=n). Each outcome x_i has a known probability p_i so that the probability of achieving an outcome of x_i or worse is written $F(x_i)=\sum_{j=1}^{i} p_j$. The RDEU functional is of the form:

$$V(\{\mathbf{x},\mathbf{p}\})=\sum_{i=1}^{n} U(x_i)h_i(\mathbf{p}) \tag{2}$$

where,

$$h_i(\mathbf{p})=q(F(x_i))-q(F(x_{i-1})). \tag{3}$$

The function q is monotonically increasing with $q(0)$=0 and $q(1)$=1.[10] Note that $h_i(\mathbf{p})$ is not a simple transformation of p_i, but is instead a function of the cumulative distribution at x_i and $x_{i\text{-}1}$.[11] The transformation function $q(\cdot)$ is where the RDEU model generalizes EU analysis. If $q(F)$=F, then the RDEU is equivalent to EU. Quiggin (1982) proposed an S-shaped function that was generalized by Tversky and Kahneman (1992) to take the form

$$q(F) = F^{\gamma} \Big/ \left(F^{\gamma} + (1-F)^{\gamma} \right)^{1/\gamma} \tag{4}$$

For $0<\gamma<1$, this specification gives more weight to the worst and best events than would be implied by the probabilities of such events, i.e. $h_i(\mathbf{p})>p_i$ for i close to 1 or n. Others (e.g., Chew, Karni, and Safra 1987) have assumed that q would be concave, which is a natural extension of the notion of risk aversion to the RDEU model. Although the RDEU adds significant nonlinear complexity to the specification of an individual's preferences, it can involve as little as only one additional parameter compared to an EU model

RDEU can explain much of the evidence that has been generated in experimental studies that conflict with the EU and SEU models.[12] As noted above, the nonlinear weights on objectively known probabilities that are observed in Allais-type behavior are easily captured by the RDEU. To explain Ellsberg-type behavior in conditions of ambiguity, in the RDEU "the decision weights are interpreted as non-additive subjective probabilities" (Quiggin 1993, p. 72; see also Epstein, 1999). Hence, this single mathematical formulation can be used to represent choices arising from nonlinear weighting of extreme events or ambiguity aversion. Still, because of its close similarities to the EU model, many of the standard results can be directly applied even if decision makers are thought to be RDEU maximizers (e.g., Quiggin 1991).

The function in (4) allows for lower and upper subadditivity.[13] It takes an inverse S-shape for values of γ between 0.27 and 1, and the parameter may be different for risks involving gains and losses. Bleichrodt and Pinto (2000) simply state that lower subadditivity, for example, means a lower interval has more impact on a decision maker than an intermediate range of probabilities. Still, it may well be that for each individual in a sample, the weighting function in equation (4) varies: there is not one weighting function that best approximates relationships for everyone in the sample. Gonzalez and Wu (1999) suggest a two-parameter specification for the probability weighting function, and also suggest a non-parametric approach, making no assumptions about the form. The latter would allow each individual to exhibit their own weighting function. Their evidence supports the two-parameter weighting function, and for their subjects the report an inverse S-shape appears “to be a regularity that holds both at the aggregate level and for individual subjects” [Gonzalez and Wu, 1999, p. 149].

Like conventional EU models, the RDEU has been and is being tested in the laboratory. A test of the RDEU is whether co-monotonic independence holds, which allows violations of independence under some circumstances. Tests of the RDEU versus EU have resulted in mixed evidence (e.g. Wakker, Erev and Weber 1994). Oliver (2003) tests the performance of the RDEU versus the EU on staff recruited from a health care facility in London. The exact experimental procedures are not well explained in the paper, but appear consistent with classifying this study as a typical laboratory experiment. Subjects are presented with health-care outcomes, treatment options, and life durations, and data are used to test the Allais paradox or violations of independence. The author concludes that when one option is certain, the RDEU explains violations of the EU. However, when all options involve uncertainty, he concludes there is no evidence supporting the RDEU as a better descriptive theory than the EU (or vice versa).

Maxmin Expected Utility (MMEU).

The second model that we highlight here is the Maxmin Expected Utility (MMEU) model developed by Gilboa and Schmeidler (1989).[14] MMEU, was developed to address the Ellsberg paradox and it formalizes the criterion of Wald (1950) and generalizes the axiomatic work of Arrow and Hurwicz (1972). Like the SEU model, MMEU does not make immediate assumptions of exogenously given probabilities, but assumes that individuals derive probabilities based on their personal experience. The MMEU differs from the SEU in the presentation of the independence axiom and through the introduction of an axiom of uncertainty aversion. As a result of these subtle differences, while SEU-rational agents would act as if they possess a unique probability distribution over the set of outcomes, **p**, MMEU-rational agents make choices based on a non-unique set of probability distributions, say, $\mathbf{C}=\{\mathbf{p}\}$. For example, in a two-outcome problem, the MMEU maximizer may believe that the probability of success is between 30 and 40 percent, but he or she is unable (or unwilling) to place a second-order probability ranking over this range of probabilities.

A decision maker following the MMEU axioms, including uncertainty aversion, would maximize the minimum expected utility. In other words, if utility is a function of choices **z** and the state x and the individual is ambiguous over a set of probability distributions, **C**, then **z** is preferred to **z'** if

$$\min_{\mathbf{p}\in\mathbf{C}} E_{\mathbf{p}}U(\mathbf{z},x) \geq \min_{\mathbf{p}\in\mathbf{C}} E_{\mathbf{p}}U(\mathbf{z}',x) \tag{5}$$

where $E_{\mathbf{p}}$ is the expectation based on the probability distribution $\mathbf{p}\in\mathbf{C}$. At one extreme, if **C** is composed of a single probability distribution then the MMEU criterion coincides with SEU theory. At the other extreme, if the decision maker is entirely ignorant of the possible probability distributions, then the criterion in this case coincides with Wald's maximin criterion.

The RDEU and MMEU models for choice under uncertainty are two leading contenders in the search for improved frameworks for choice under uncertainty. Still, Harless and Camerer's 1994 conclusion appears to still hold true today – there are no clear winners; some theories explain little behavior while others cannot discriminate between behavior consistent with the model and purely random choices. Some models have strong axiomatic foundations while others are essentially ad hoc. Loomes's (2006b) recent work echoes these concerns about the lack of a general model. Despite this unresolved (or unresolvable) state of affairs, non-EU models are already being found useful in applied economic analysis. For environmental and resource economists, we see important implications for the theory that underlies *our* work, in the issues that *we* study, and in the methods that *we* employ.

Non-EU analysis in Resource Economics: The theory we use

Like it or not, benefit-cost analysis (BCA) is one of the most basic tools *we* employ. With conditions of uncertainty, there is still some debate about how BCA should proceed (Graham 1981, 1992; Ready 1993, 1995; Chavas and Mullarkey, 2002). However, it is seems there is growing consensus that, using the definitions of Hammond (1981), it is appropriate to use an ex ante measure, i.e. based on individual's willingness to pay in an uncertain world, as opposed to an expected ex post measure in welfare analysis, or at least BCA.

Although it appears that the profession has largely accepted this ex ante perspective, the question of which ex ante welfare measure is appropriate has not been resolved formally, and may in fact be unresolvable. Hammond (1981) spells out the contrasting points of view. On the one hand, a respect for "consumer sovereignty" favors an ex ante approach. As summarized by Krupnick, Markandya, and Nickell (1993, p. 1274, emphasis added), "since damages are estimated based on individuals' willingness to pay to avoid risks, it is more appropriate to estimate damages based on lay perceptions of risk since *their* willingness to pay is based on *their* perceptions."

On the other hand, the fact that "individuals misperceive the probabilities of certain events," suggests that perhaps a benevolent (paternalistic) planner might be in a better position to advance the interests of the misguided public. From this perspective the analytical objective would be to identify the expected impacts of policies after the fact, i.e., the expected ex post value. The alternative perspective is described by Loomes (2006a):

> "the best estimate of the loss of wellbeing entailed by becoming diabetic, or suffering non-fatal injuries in a road accident, or having a new housing development built on nearby farmland can be obtained by measuring what life is actually like for people who have those experiences compared with people who have not."

While not recommending an ex post approach, Crocker, Forster, and Shogren (1991, p. 6) present the case of an environmental danger that is not sufficiently appreciated by the public, "Does the regulator ban the environmental danger or allow individuals to use their own discretion? The dilemma is to balance the tradeoff between preserving individual freedom of choice and maintaining public safety."

As discussed above, individuals sometimes place more or less weight on risks than would be merited based on the best available estimates of the probabilities. As highlighted in Table 1, this can lead to conflicts if policy-makers side with the experts and the general public's perceptions fall at the other end of the spectrum. Consider the control of risks associated with nuclear waste disposal (again, as included in Table 2). Two options might be available: Plan A with very small risks and high costs, and Plan B with lower risks and extraordinarily high cost. Based on expert assessment of the risks, standard benefit-cost analysis may favor a Plan A, exposing the population surrounding the transportation route to very small risks of severe consequences. The public, however, because of ambiguity surrounding the risks (for lack of information) or because of tendency to overweight low-probability high-consequence events, may be willing to pay extraordinary costs of Plan B to reduce risks even further. Conflict will arise if Plan A is adopted despite the fact that, based on best available evidence, Plan B does not pass a benefit costs test and would, on average, be a mis-allocation of resources.

Policy analysts, therefore, have a clear choice. Respect consumer sovereignty, or watch out for the best interests of the public, ex post. The notion of consumer sovereignty is a deeply embedded tradition in economics, and it is part of the Pareto criterion. Even in the general welfare economics tradition of Kaldor (1939) and Hicks (1939), the notion that a government's choices should be built on the preferences of individuals is widely accepted. Yet if non-EU preferences prevail, there will be a conflict between what is preferred by individuals and what is, ex post and on average, in their best interests. It is not even clear how choices would be made by elected officials; if most relevant uncertainty is resolved within the electoral cycle, then respecting consumer preferences would be politically wise, while if uncertainty is resolved outside the political cycle, then the expected ex post perspective is politically preferred.

The answer to this question also has important implications for how we carry out non-market valuation (often used to obtain the benefits side in BCA) and policy analysis. Stated preference valuation methods would typically capture the ex ante preferences of individuals. If the individual's preferences are RDEU as defined above, a policy z that affects the vector of goods and services $\mathbf{x}$ would yield ex ante welfare $V(z) = \sum_{i=1}^{n} U(\mathbf{x}_i(z)) h_i(\mathbf{p})$. The monetary value found in stated preference methods would be a monotonic function of $V(z)$. On the other hand, as noted in the Loomes quote above, revealed-preference data are often observed after the state of nature is revealed, so that the valuation effort is related to the expected ex post utility function, $U(\mathbf{x}_i(z))$.[15]

Prior to beginning a valuation exercise, therefore, it is important to resolve the question of whether values should capture ex ante preferences or expected ex post outcomes. Given that there are legitimate questions about whether an ex ante or expected ex post measures are appropriate, the most flexible approach would be for applied benefit-cost analysts to carry out both, and to pay specific attention to discrepancies that arise because of non-EU preferences. If such complete analysis is not feasible, then the perspective maintained should be stated explicitly.

A Role for non-EU analysis: Some issues *we* study

As highlighted in Table 2, there are many problems that *we* study where conditions are ripe for EU axiom violations to occur. First, as noted by Toman (1998), non-SEU models might be appropriate when evaluating the climate-change policies. Putting it more strongly, Geoffrey Heal and Bengt Kriström (1993) state:

> Even if we knew exactly what the climate would be in 2050, we still would face major economic uncertainties because we currently do not know how altered climate states map into human welfare. (p. 4).

Simply recognizing uncertainty and incorporating it into a standard climate change model makes it optimal to more aggressively seek to reduce warming (Ayong Le Kama and Schubert 2004).[16] However, the degree of ambiguity about the consequences of climate change is substantial among climate change scientists, prompting Cameron (2005) to incorporate ambiguity into her modeling of behavior related to climate change. Woodward and Bishop (1997) also explicitly address ambiguity, considering the case of a social planner who uses the maximin framework of Arrow and Hurwicz (1972). Whether individuals place nonlinear weights on probabilities or are averse to ambiguity, the general result would be the same – compared to with EU preferences, such individuals would place more weight on the most severe consequences of global warming. These non-SEU frameworks, therefore, provide a justification for more aggressive climate policies, akin to policies that follow from the *precautionary principle*.

Another issue for which non-EU preferences could play a role is in assessing the risks of extreme natural events such as floods, droughts, some health risks, and other catastrophes. Several studies indicate that markets in insurance are not working correctly in these areas (e.g. Chivers and Flores 2002; Kunreuther and Pauly 2004). As a specific and probably poorly known example, White (2001) suggests that options market for catastrophes have not worked as well as had been hoped. The recent analysis of nuclear wastes and responses to extremely low probabilities as risks (Riddel and Shaw 2006) sheds light on how people react to catastrophic/low probability events.

As a final issue, consider all of those risks that might be characterized as being endogenous to the individual. Several authors have argued that risks are actually endogenous because individuals can self-protect or mitigate to reduce the severity of an outcome that has negative consequences (e.g. Shogren and Crocker 1991). As Heal and Kriström (1993) note, climate change risks are "generated by our own activities," at least to some extent. Using another example from the literature, the probability of lightening striking a home is exogenously determined, but the probability of the house burning down is not. Yet, economists are generally more interested in health outcomes (injuries and death) than the lightning event itself. If that is the case, then one would not wish to use objectively-determined probabilities in an EU model, though it may be possible to map from the probability of lightening to the probability of a health outcome. Quiggin (1992) added considerable detail to the original framework that Shogren and Crocker (1991) laid out, and in Quiggin (2002) shows the applicability of the general concepts to a state-contingent approach, which is then applicable to non-expected utility formulations. Ideas how endogenous risk can be incorporated into econometric models are presented below.

A Role for non-EU analysis: The methods *we* employ

From the tools of non-market valuation methods to dynamic optimization, there is a need for determining how non-EU preferences should be incorporated into the tools of our trade. As an example, Viscusi (1997) reports on failure in rational Bayesian learning and alarmist reactions to health risks associated with air pollution. Basic work on risk preferences related to environmental and resource issues needs to be done, and we should be trying to find the shape of the PWF if we think it might be non-linear. Such shapes have been shown to depend on the outcome domain (Bleichrodt and Pinto 2000).

For example, Heal and Kriström (1993) noted in 1993 that "…the question of risk aversion [had] not been studied in the context of climate change." [p. 14]. To our knowledge, it still hasn't been. Other examples for resource economists follow below.

Valuation

Though it sometimes offends the sensibility of theoretical micro-economists, preferences estimated using stated choice data are often used to estimate values. This is more common with situations where certainty is assumed than when risk/uncertainty is allowed, but there are estimates of welfare under conditions of uncertainty that exist, mostly based on the EU framework. If actual preferences are not consistent with EU framework, then the resulting estimated values will be biased. In his generally critical review of non-EU models, McFadden (1999, p. 97) notes that the standard models "does not apply universally, or even regularly, to choices made in non-market contexts." He goes on to note, "Nowhere has this been more evident than in economists' attempts to value non-use public goods, such as endangered species or wilderness areas." Yet most applications that we know of in environmental economics stick closely to the basic EU framework's assumptions in analyzing choices (see Shaw, Jakus, and

Riddel 2005, for a summary). Recent exploration of non-EU preferences, and particularly, allowance for choice under ambiguity, can be found in the food-safety and health risk literature (see Kivi and Shogren 2005; Shaw, Nayga, and Silva 2006).

As noted above, econometric methods can be used to test for the structure of the expected utility function, allowing determination of whether probabilities enter linearly or not (Riddle and Shaw (2006; 2003) and Cameron (2005)). The resulting ex ante welfare measures in these studies takes a form that is different than the usual option price equation that would be derived in the linear-in-probabilities form of Graham (1981). If nonlinear probabilities or ambiguity aversion affects individuals' preferences, then these alternative forms should be recognized in the valuation models that are employed. It is relatively straight-forward to discern whether subjects feel uncertain about the risks that they face, and some survey questionnaires even induce ambiguity, so it is likely impossible to avoid.

Another issue in valuation involving uncertainty is when respondents to contingent valuation questions say they don't really know the answers to questions, or at least they are uncertain about the exact magnitude of the value of something. In some cases, from the individual's point of view, this may overlap with actual risk, but this need not be the case. An interesting approach is to allow for "fuzzy" logic or math in the derivation of the resulting values in this context. Recently van Kooten, Kremar, and Bulte (2001) and Sun and van Kooten (2005) have demonstrated that fuzzy set theory may be used in lieu of the other approaches to dealing with uncertainty values. Loomes (2006a), on the other hand, interprets all contingent valuation through a skeptical psychological lens that questions the ability of individuals to provide meaningful translation of anticipated events into monetary measures of willingness to pay. What all of these approaches have in common is that they relax, at least to some degree, the assumption that individuals have EU type preferences.

It is also possible that the shortcomings of the EU model might be behind the frequently observed differences behind estimates of willingness to pay and willingness to accept that are obtained in a context of risk. Horowitz and McConnell (2003) have argued that the magnitude of the WTA/WTP ratios found in the literature are "implausibly high" and "reject the claim that observed WTA/WTP ratios are consistent with a standard neoclassical model." Empirical evidence that the WTA-WTP disparity may be related to non-EU preferences is mixed. With a small sample of respondents, Brown (2004) found some evidence that ambiguity may be related to the divergence, but did not formally test this. Eisenberger and Weber (1995) use a simple event lottery, allowing experimental subjects to specify their maximum WTP to participate and minimum WTA to not participate. They do not find evidence of an interaction between ambiguity and the WTA/WTP ratios, but this is but one of the first experiments to explore ambiguity's role in this regard. Understanding why WTA and WTP measures differ so much has important practical implications, and the search for the answer certainly cannot exclude the role that non-EU preferences might have in the answer.

Our final point regarding valuation is that econometric methods could be adapted to the problem of endogenous risk. Statistical measures are available to test for endogeneity of risks and this can be important in valuation of risk changes. If risk levels that are faced by individuals are truly endogenous, then it would not be appropriate to treat them as exogenous in econometric models, regardless of whether they are to be used within an EU or non-EU framework. Tests for exogeneity are quite common in econometrics, yet this type of test is not common when studying environmental risks. The key econometric issue, presuming that perceived risks are central in explaining behavior, is whether these risks vary independently from other variables in the model. Engle, Hendry, and Richard (1983) define conditions where a set of variables is weakly exogenous in the context of a full model: this boils down to whether the full model is equally efficient by using a conditional or full, joint distribution for the set of variables. For example, suppose that we are modeling drinking water in the presence of health risks, as a function of perceived health risk, income, and other variables. One can imagine a possible relationship between the formation of risk perceptions and income, as income may also relate to education, concerns about paying for health care, etc. It would seem prudent to test for weak exogeneity.

Suppose that tests show that risks are endogenous. If so, there are grave concerns about moral hazard issues and the role of markets or insurance (Heal and Kriström, 1993); uncovering estimates of risk may be difficult, involving elicitation of perceived and subjective risks. What then? We would recommend that researchers explore a model where perceived risks are potentially dependent on key variables, and then feed predicted values of the risk back into the behavioral model. Those variables might be age, gender, and factors that probably also determine whether a person purchases health insurance (e.g. income, family size, past health history). Such modeling of perceived risks is certainly not new, or beyond researchers' capabilities: Viscusi and Evans (1998) jointly estimated preferences for product safety decisions and risk perceptions, and Riddel and Shaw (2006) feed predicted subjective risks into their behavioral model of the decision to move away from routes carrying high-level radioactive wastes.

What has not been done, to our knowledge, is then to further explore whether the predicted values of the modeled risks are consistent with EU- or SEU-type behavior. This is going to be difficult, as conclusions about the validity of EU or SEU assumptions rely on convincing modeling of the behaviors over the entire region of subjective risks, not just some portion. In other words, we might elicit subjective risks only in some portion of a range of probabilities, but find that our endogenous risk model predicts quite poorly outside this range. However, for policy purposes *we* might be on safe ground, as marginal risk changes from programs to reduce risk are much more likely than huge changes in risk.

Optimal Resource Management

Another area in which non-EU preferences are beginning to be applied is in the analysis of optimal resource management. Analysis of optimal resource management typically makes use of the dynamic optimization frameworks of optimal control and dynamic programming or their stochastic counterparts. Building on Kreps and Porteus (1978), Epstein and Zin (1989) provide a tractable framework for carrying out dynamic optimization without assuming either additively separable welfare across time or expected utility preferences outcomes. Their model, therefore, could accommodate preferences that are nonlinear in probabilities. The Epstein and Zin framework might be used in applied analysis, but there are relatively few such applications. In the area of natural resource management, Knapp and Olson (1996), Peltola and Knapp (2001), and Epaulard and Pommeret (2003) apply the Epstein and Zin formulation, though in each case the attention is primarily on separating time preferences from risk attitudes without relaxing the expected utility formulation.

Another thread of interest is the use of *robust control* methods as has been carried out by Roseta-Palma and Xepapadeas (2004). Robust control is an optimization approach that is widely employed in engineering and optimizes relative to the worst-case scenario. This is particularly appropriate when designing, for example, a bridge, since “getting it right” almost all the time is more important than “getting it right” on average. Hansen and Sargent (e.g. Hansen et al. 2006) have begun to advocate the use of robust control in macroeconomics, and motivate it as an implementation of the MMEU of Gilboa and Schmeidler (1989). Roseta-Palma and Xepapadeas have applied the same logic to the resource management problems in the areas of water (2004) and fisheries (Xepapadeas and Roseta-Palma, 2003). The jury is still out on whether robust control is appropriate in an economic context (Sims, 2001). However, if it is applicable anywhere, then we believe that it is certainly so in many problems of environmental and resource management.

Summary and some caveats

In this paper we have summarized some of the evidence pertaining to the axioms of the expected utility and subjective expected utility hypotheses, making this accessible to environmental economists who might not be looking at the more general literature on risk and uncertainty and perhaps having difficulty making the connections. Consensus that the EU and SEU models fail has not yet been reached, but it appears that many decision and uncertainty theorists are leaning away from them (Fishburn 1988; Starmer 2000). Should *we* (environmental and resource economists) care one way or the other? We believe so.

As Table 2 makes clear, for many of the problems *we* study the conditions are ripe for departures from the EU framework. Low probability-high consequence events are prevalent throughout environmental and natural resource economics. *We* often deal with problems in which scientific understanding is weak and the public may face significant ambiguity. So, we believe, resource economists should take note of the new empirical and theoretical approaches that are being developed to handle departures from the EU framework.

However, we do not want to give the impression that this is the only or most important conceptual challenge that our discipline faces. As reviewers and an editor reminded us, it is important to remember two things. First, there are many failed assumptions that can be reassessed, and where having done so, behavior can be better explained; usually these assumptions have little or nothing to do with fundamental tenants of the EU. The question of rationality lurks behind all modeling of behavior, and Cherry, Crocker and Shogren (2003) explore whether rationality with market choices spills over to choices involving non-market situations; they find evidence that it does. They do not find evidence that preferences for lotteries change because of the context. A large body of research has been devoted to the stability of preferences, much of it within the rubric of certainty. Richard Thaler (1981) has contributed to this vast literature, finding many inconsistencies in preferences. Jack Knetsch, as another example, demonstrated many times that preferences under certainty may differ depending on whether the individual faces losses or gains (see Knetsch 1989; his discussion of some of these in Knetsch, 1995). Allowing for such differences can help explain behaviors otherwise puzzling. Another example of inconsistencies in preferences relates to discounting: people may change their private rate of discount depending on the length of time being considered or other contexts (Lowenstein and Prelec 1992). Finally, Loomes (2006b) provides a powerful critique of almost all models of rationality (including his own) arguing that apparent violations of the EU can be explained simply by the inability to make fine distinctions in multi-attribute choices.

Yet another example of an invalid assumption that carries over to the EU and any other model is that there are no transactions costs in making decisions. For example, there are many instances in which individuals fail to purchase insurance when the basic EU model would suggest they should. Such discrepancies can often be reconciled by assuming that there are transactions costs that drive a wedge between EU maximization and observed behavior. Kunreuther and Pauly (2004) consider the following paradox: if the probability of an earthquake next year is 1/250 or greater, and the ratio of the insurance premium to the damage is less than or equal to 1/250, then one would expect to see risk neutral or averse people making purchases of this actuarially fair insurance. They do not often see such purchases being made. They maintain that individuals do try to maximize expected utility, but they face the cost of discovering the true probability of these events, and this creates a block to the purchase. Like imperfect information, the cost of obtaining and processing information may be the root cause of inconsistency with EU predictions in actual market settings.

The second reminder relates to the fact that EU may work well in some situations, and that non-EU models may also fail to adequately predict behavior. If probabilities are well-understood, perhaps when they fall in the middle range (e.g., a fair coin toss), and the magnitude of difference between losses and gains is quite small in some gamble, the EU may adequately predict what choices people will make. Some have argued that if financial incentives (e.g. scoring approaches) are used on survey or experimental subjects, they will provide subjective risk assessments that are consistent with those based on so-called experts, and once that has been achieved, the EU is the acceptable model to apply to the data.

Markets that function well to reward and penalize those making gambles have been shown to reduce violations of the EU (e.g. Evans 1997). This suggests a role in some risk allocation contexts for options contracts. However, Dow and Werlang (1992) show that ambiguity can diminish the actual trade in such risk-hedging instruments. And, of course, the focus of this paper is on environmental and resource issues, which are largely characterized by non-market goods.

Neither of the non-EU models we favor predict all the choice patterns that are frequently observed. For example, Quiggin's RDEU framework implies that individuals who overweight low probability outcomes are pessimists who will avoid these lotteries, yet Battalio et al. (1990)

find no support for the RDEU's predicted behaviors. Further, RDEU is designed to maintain transitivity, but Loomes (2006b) notes that problems in which violations of transitivity are common. Similarly, Machina's (1982) fanning out model precludes behavior consistent with "fanning in," but some people exhibit that.

The problem with some non-EU alternatives rests with the fact that the EU is usually a special case. Much of the behavior inconsistent with the EU, therefore, will also be inconsistent with its non-EU parents.

Final Thoughts

For environmental policies the problem of ambiguity is often central and our normative analysis should appreciate its importance. We offered the example of climate change, and as Faucheux and Froger (1995, p. 31) point out, "global environmental problems have no historical precedents…. This means that the information on which decisions are made is most often a non-probabilistic kind." Therefore, as economists enter into these policy debates, we should keep in mind that our recommendations are only as correct as the axiomatic foundation of the underlying decision model (Woodward and Bishop 1997).

In 1738 Daniel Bernoulli solved the St. Petersburg paradox by inventing the principle of expected utility. Today we look with puzzlement at the response of Daniel's cousin Nicholas Bernoulli who continued to look in vain for a single fair price for the game based on its expected value. After retelling this story, Harless, and Camerer (1994) ask, "Might future economists find it [equally] peculiar that twentieth century economists held firmly to EU in the face of the Allais paradox and other violations?" (p. 1284). There may be no general theory of decision making under uncertainty. Graham Loomes (2006b) recently states "if [his] model of the data generating process is correct, there can be no descriptively adequate general theory of risky choice which is rational..." [p. 2]. Still, some otherwise puzzling outcomes can be explained by alternatives to the EU models, and we believe that environmental and resource economists should begin to pay attention to these alternative models.

Endnotes

[1] A reviewer notes that several studies suggest possible inconsistencies with conventional EU models, but that in many of these studies the individual's behavior is in fact consistent with the state-dependent version of the EU, or the EU considered with transactions costs. Examples are tax evasion (Bernasconi 1998); criminal behavior (Neilson and Winter 1997), financial markets (Dow and Werlang ,1992; insurance against catastrophic events (Kunreuther and Pauly 2004), and consumer willingness to partake in precautionary saving (Leland 1968). In each instance, reasonable explanations for behavior that is consistent with the EU can be found by assuming state-dependence or adding transactions costs.

[2] See Anand (1993) for a review.

[3] Wakker and Deneffe (1996) conclude that the "trade-off" approach has distinct advantages over the other methods (the certainty-equivalent (CE), direct scaling, and probability-equivalent methods).

[4] Results are also based on an experiment involving monetary gambles. It is also worth a reminder here that all of the subjects used by Wakker and Deneffe were, as they put it, "well acquainted with probabilities and expectations." [p. 1135]. The subjects used in our experiment certainly appear to be not so well-acquainted.

[5] To be fair here, we note that Karni and Safra (1995) argue that experimental economics cannot be a decisive tool to use when it comes to eliciting subjective risks. They offer a proof that subjective risk estimates obtained when the subjects have a stake in outcomes pertaining to total wealth are simply impossible to elicit. Their strong-sounding conclusion is softened when outcomes are couched as losses, as many environmental economists might express environmental damage, or gains/improvements, as many might express the environmental improvements that have arisen because of the Clean Air Act and the like. Karni and Safra's conclusion therefore does not negate all of the conclusions that flow from studies of subjective risks. Still, it would be more convincing if evidence of the failures of the EU framework came not from the laboratory, but from observed behavior of economic agents.

[6] Anand (1993) provides a good overview of the experimental work, breaking it down significantly more than we have here. Harless and Camerer (1994) survey and synthesize the experimental evidence related to Allais-type behavior, and Camerer and Weber (1992) review the studies that have addressed Ellsberg-type behavior and ambiguity. Shaw, Riddel and Jakus (2005) survey the literature with a focus on valuation.

[7] Skeptics might say that if the experiment just incorporates incentives for getting risks in line with the experts, then the PWF for an individual will become linear, i.e. people are EU

maximizers. For example, Harrison and Rutström (2006) pay subjects for correctly ranking mortality risks the same way that the data from sources like the Center for Disease Control do, thereby showing that discrepancies between the CDC rankings and the subjects are small. This again supports the notion that EU modeling is more plausible when risks (here for other, anonymous people) are well-understood. We are not sure of the usefulness of this experimental result on ranking risks: for one, their incentive scheme might do little to change risk rankings when the mortality risks are personal rather than for other people. The PWF for environmental risks is (to our knowledge) an unexplored empirical question.

[8] Readers are referred to Kelsey and Quiggin (1992), Camerer and Weber (1992), Quiggin (1993) and Harless and Camerer (1994) for much broader reviews of the literature and perspectives on the relative utility of different models.

[9] The RDEU model has, however, also been applied to conditions of ambiguity by Segal (1987), who uses the RDEU to apply to second-order probability distributions. Similarly, the model is quite similar to the models of Schmeidler (1989) and Gilboa (1987) that axiomatize a model that uses Choquet integration when decision makers are faced with ambiguity.

[10] The assumption that $q(1)=1$ is relaxed by Dow and Werlang (1992) to address the issue of ambiguity.

[11] This property is what guarantees that the stochastic dominance is satisfied and is roughly equivalent to Choquet integration that has been axiomatically derived by Schmeidler (1989).

[12] Harless and Camerer (1994) point out, however, that compared to other models the RDEU framework is not very discriminating so that tests of the model have relatively low power.

[13] The weighting function satisfies subadditivity if there exist constants $\varepsilon \geq 0$ and $\varepsilon' \geq 0$ such that $w(q) \geq w(p+q) - w(p)$ whenever $p + q \leq 1 - \varepsilon$. Here q is a lower value for a probability and p is a higher one. A similar expression pertains to upper subadditivity. See Figure 2 in the paper by Tversky and Wakker (1995). Without reproducing that graph here, the lower subadditivity condition basically implies a very steep portion of the weighting function near zero, with a flattening in the middle portion.

[14] A similar framework has also been proposed by Kelsey (1993).

[15] Notice that this also suggests a reason why convergent validity may not be achieved.

[16] Howarth (1997) notes that there is an additional dimension to uncertainty: not only are we uncertain about the consequences of climate change, but also about the preferences of future generations.

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Table 1: Uncertainty Conflicts between the General Public and Policy Makers/Government

		Policy Makers/Experts	
		Low risk assessment	*High risk assessment*
General Public	*Low risk assessment*	***Little conflict*** Low policy investment and public mitigation	***High conflict*** Policy makers need to communicate risks/change perceptions. Public sees regulation/government expenditures as unnecessary and wasteful; public demand for risk-reducing measures is low
	High risk assessment	***High conflict.*** Public engaged in unnecessary expenditures/effort on mitigation; public sees policy makers/government as shirking responsibility; demand for risk-related goods too low	***Little conflict*** High policy investment and public mitigation

Table 2: Examples of Risk Situations of Interest to Resource Economists*

Situation (type of risk)	**Relevant probabilities are small or very small**	**Ambiguity is important**
Agricultural Crop Failure (Financial)		
Airline Safety (Mortality)	✓	
Arsenic In Drinking Water (Mortality/Morbidity)	✓	✓
Climate change risks		✓
Extinction Of Species (Species Mortality)	✓	✓
Financial risks in resource management		
Flood/Hurricane Risks (Financial, Mortality)		✓
Future Resource Prices (Financial)		
Giardosis (Mainly Morbidity)	✓	
Human Health And Air Pollution(Mortality/Morbidity)	✓	
Invasive Species (Harm To Native Species/Habitat)	✓	✓
Non-point source pollution (emissions and damages)		
Oil Exploration (Lower Prices/Financial)	✓	
Radioactive/Nuclear Waste(Mortality/Morbidity)	✓	✓
Risks of enforcement of pollution non-compliance	✓	
Success in recreational fishing or hunting		

* The check marks indicate our personal assessments of risks involved and the public perceptions of those risks. This is not intended to be conclusive, but indicative of the variety of problems studied in the area of environmental and resource economics.

Intra-Regional Amenities, Wages, and Home Prices: The Role of Forests in the Southwest

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Intra-Regional Amenities, Wages, and Home Prices: The Role of Forests in the Southwest

Abstract:

Forests provide non-market goods and services that people are implicitly willing to pay for through hedonic housing and labor markets. But it is unclear if compensating differentials arise in these markets at the regional level. The answer to this empirical question is addressed in a study of the Southwest United States, composed of Arizona and New Mexico. Hedonic regressions of housing prices and wages using Census and geographic information systems data show that U.S. Forest Service area carries a marginal implicit price of between $27 and $36 per square mile annually. Surface water area is priced between $4 and $16 per square mile annually. The existence of significant compensating differentials at the regional level suggests that care must be taken when applying the travel cost method to value regionally-delineated characteristics. (JEL: Q23, R14, R23, J31)

Keywords: Compensating differentials, hedonic prices, forests, amenities

1 Introduction

As traditional resource extraction has become relatively less important in regional economies of the American West, the amenity values of forests and other natural features have received more attention. Instead of seeking out areas that, for example, hold the promise of a large paycheck from logging or mining jobs, migrants may seek out areas that offer a large so-called second paycheck derived from the value of natural landscape. This economic behavior can have observable effects in the markets for housing and labor.[1]

As argued by some authors (e.g., Niemi et al. 1999), a "second paycheck" may be derived from forests because these areas contain many of the characteristics that people value and make economic decisions over, such as abundant recreation opportunities, open space, and the availability of wilderness experiences. Further, forests provide several non-market ecosystem services, such as watershed and airshed management, soil quality protection, and habitat protection. Individuals who value these characteristics may be willing to pay, through compensating differentials in labor and housing markets, for access to these non-market goods and services. A portion of the first paycheck (i.e., money income) may be forfeited in favor of a larger second paycheck.

Compensating differentials have been shown to exist in labor and housing markets for large-scale amenities (e.g. climate) at the inter-regional level (Roback 1982; Hoehn et al. 1987; Blomquist et al. 1988; Costa and Kahn 2003), and in housing markets for local or neighborhood amenities (Ridker and Henning 1967; Cheshire and Sheppard 1995; Acharya and Bennett 2001). Studies that include forest resources have also shown that proximity to forests is capitalized into local housing prices (Shultz and King 2001; Kim and Johnson 2002; Kim and Wells 2005) and wage differences between cities (Schmidt and Courant 2006).

Lacking in the literature, however, is an understanding of how forests (and other natural characteristics) impact economic behavior within a particular region. Whether or not forest amenities are important in the housing or labor markets (or both) within a region remains an empirical question.[2]

Compensating differentials imply that the supply of forest amenities at the regional level is policy relevant. National forests in the U.S., under the management of the U.S. Department of Agriculture's Forest Service, undergoes a periodic regional planning process.[3] In the Forest Service's Southwestern region, this planning process involves regional teams formed to ensure coordinated revisions of all individual forest plans.[4] Knowledge of the importance of forest resources at the regional level would aid in determining the economic impact of regional forest policy decisions and how management decisions conform to the public's preferences.

The purpose of this paper is to examine compensating differentials for forest amenities in housing and labor markets in the Southwest United States, composed of Arizona and New Mexico. This region provides an excellent opportunity to answer the research questions asked in this paper. The region is characterized by wide variation in the size, type, and location of forest characteristics and their proximity to large population centers. A unique geographic information system (GIS) data set from the U.S. Forest Service Southwestern region allows these characteristics to be precisely measured. Hedonic regressions are estimated for housing prices and wages using publicly available Census data, GIS calculations of landscape features, and data enumerating outdoor recreation sites. An empirical framework pioneered by Roback (1982) and developed by Hoehn et al. (1987) is used to determine the existence of hedonic premiums for forest amenities in housing and labor markets and the implicit prices faced by Southwestern region residents for these amenities.

2 A theory of regional hedonic wages and housing prices

Forest characteristics have been unambiguously shown to affect labor and housing markets in a variety of empirical settings. Among these are studies of the wage and housing price effects of interregional differences in forest access (Schmidt and Courant 2006) and local forest access (Kim and Wells 2005). The primary empirical question in this paper is whether forest characteristics that vary regionally are in fact amenities that significantly affect regional housing price and wage differentials.

The basic idea behind compensating differentials in housing and labor markets is that individuals purchase a bundle of characteristics tied to a heterogeneous good in the marketplace (see Rosen 1974). In the context of regional forests, the bundle of characteristics is tied to where one lives and works. Someone living in Mora County, New Mexico (population about 5,100) has access to a different set of forest, outdoor recreation, and other characteristics than someone living in downtown Phoenix, Arizona. Further, the labor market opportunities are unique to each place. These differences define the bundles that people make decisions over and may give rise to compensating differentials.

The model described here is most similar to that in Roback (1982). Several theoretical advances have been made since this model was developed (e.g., Hoehn et al. 1987), but the Roback model is more appropriate to this application because the study area does not have endogenous geographic borders. Although the model presented here does not differ substantially from that in Roback (1982), the key points are presented to describe the conditions under which compensating differentials will arise within a region and to introduce the notation with respect to forest characteristics.

The simplest model of compensating differentials only considers individual housing location decisions (ignoring firms and endogenous income for now). Suppose individual utility is a function of numeraire good consumption, X, land area purchased for housing, L at price p, and a forest amenity f.[5] The consumer's problem is:

$$\max \;\; U = U(X, L, f) \;\; s.t. \;\; M = X + pL, \tag{1}$$

where M is income. The choice variables are X and L, with the choice of f being implicit and determined by one's residential location. A convenient way of thinking about the consumer's problem is that they select X and L to maximize utility given the level of f where they live.

Optimization of (1), conditional on f, yields the indirect utility function, $V = V(p, M, f)$. In spatial equilibrium, utility differences are arbitraged to zero as a result of free and costless mobility. Imposing this spatial equilibrium is done by holding utility constant across $j = [1, 2, \ldots, J]$ locations, or

$$V_j = V(p_j, M, f_j) = V_0 \quad \forall j. \tag{2}$$

Spatial heterogeneity within the region is a necessary condition for the existence of hedonic premiums. This necessary condition is met if movements across locations in the region result in changes in f. To see how this heterogeneity affects the existence of hedonic premiums, totally differentiate (3) and rearrange, noting that $dV_0 = 0$:

$$\frac{V_f}{V_M} = -\frac{V_p}{V_M}\frac{dp}{df} - \frac{dM}{df}, \tag{3}$$

where the subscripts of V refer to partial derivatives. Since X is a numeraire good, the marginal effect of a change in income is equivalent to the marginal utility of X. The term on the left-hand side of (3) is the marginal rate of substitution between the forest amenity and goods consumption; this term is the implicit marginal price of the amenity.

Equation (3) implicitly defines the demand for amenities and what individuals are willing to trade to get them. But whether or not such tradeoffs are made depends on whether they are available in a market. Roback (1982) suggests that a sufficient condition to generate market tradeoffs is that the supply of land at every location is limited, and individuals cannot occupy the same space in time. Given spatial heterogeneity, the price of land must adjust across areas such that

$$V(p_1, M, f_1) = V(p_2, M, f_2) = \ldots = V(p_J, M, f_J). \tag{4}$$

For continuous variations in f, we can then observe the familiar hedonic housing price function, the price schedule as a function of the different amenity levels, or $p = p(f)$.

The second term on the right-hand side of (3) says that the total implicit marginal price people are willing to pay for amenities includes their willingness to trade income for amenity access. Is there any reason to believe that the market will offer these tradeoffs? Suppose income is wholly derived from labor market wages, so $M = w_j$ where w_j is the wage rate in area j and individuals supply one unit of labor. Following Roback (1982), the unit cost function for any firm must satisfy $C(p, w, f) = 1$. By assumption, $C_p > 0$ and $C_w > 0$. No assumption is initially made about the sign of C_f.

In equilibrium the model must satisfy,

$$V_0 = V(p_j, w_j, f_j) \quad \forall j \tag{5}$$

$$C(p_j, w_j, f_j) = 1 \quad \forall j \tag{6}$$

where V_0 is a constant. The first equation defines utility equality of individuals across all areas, while the second defines production cost equality. Totally differentiating (5) and (6), noting that $dV = dC = 0$ in equilibrium, and solving for the two unknowns $\frac{dw}{df}$ and $\frac{dp}{df}$ yields partials that are functionally equivalent to equations (4) in Roback (1982):

$$\frac{dw}{df} = \frac{V_f C_p - V_p C_f}{D} \tag{7}$$

$$\frac{dp}{df} = \frac{V_w C_f - C_w V_f}{D} \tag{8}$$

where $D = -V_w C_p + V_p C_w < 0$.

As noted in Roback (1982) and Hoehn et al. (1987), the signs of the partials in (7) and (8) depend on the sign of C_f. Characteristics that are amenable to individuals may or may not affect productivity for firms. The classic examples are highway infrastructure (cost reducing), clean air (cost increasing) , and sunshine (productivity neutral). If $C_f > 0$ (the amenity is unproductive), then $\frac{dw}{df} < 0$ and $\frac{dp}{df} \lesseqgtr 0$. $C_f < 0$ (the amenity is productive) implies $\frac{dw}{df} \lesseqgtr 0$ and $\frac{dp}{df} > 0$. In the case that amenities do not matter for firms, both partials are unambiguously signed less than zero and greater than zero, respectively.

Equations (7) and (8) indicate that wages and housing prices will vary with the level of amenities one has access to; the sign and magnitude of these variations will depend on the structural parameters of utility functions and production relationships. These partials are not equal to zero only under certain conditions: movements across the landscape result in changes in amenity levels, individuals cannot occupy the same space in time, and it is costly for individuals to access amenities in areas where they do not live.

The first condition is met by the fact that the characteristics of interest, those represented by forests, vary widely across the study region. The second condition is assumed as in Roback (1982). The final condition is the most relevant for estimating compensating differentials within a region. Some areas within the study region considered in this paper will more plausibly meet this condition than others. For example, people may live and work in the greater Phoenix area but commute across geographic areas as defined by the available data. This issue is discussed as necessary in the empirical sections below.

3 The hedonic empirical framework

The partial derivatives of (6) and (7) are what traditional hedonic studies seek to estimate. Beginning with housing price, forest amenities are but one feature that impact price. The price of a house is a function of its structural (square feet, number of bedrooms), neighborhood (crime rates, school quality), and forest amenity (forest views, recreation opportunities) characteristics. The hedonic price function specifies the price of house i as,

$$P_i = P_i(S_i, N_i, F_i) \tag{9}$$

where S, N, and F represent vectors of structural, neighborhood, and forest amenity characteristics, respectively.

The generalized form suggests that each housing unit can have unique values for each argument in the price function. It is commonly assumed that units in a given area share some characteristics, like school quality (since children in a particular school district have the same public school option). This assumption also fits well with the regional hedonic model described above. For the purposes of this paper, it is assumed that the neighborhood and forest amenity characteristics are common to all housing units in a given geographic area j, or

$$P_{ij} = P_{ij}(S_i, N_j, F_j). \tag{10}$$

Converting equation (10) into an econometric model yields

$$P_{ij} = \alpha_0 + \beta' S_i + \gamma' N_j + \delta' F_j + \upsilon_i, \tag{11}$$

where υ_i is the assumed zero-mean random error and α_0, $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$, and $\boldsymbol{\delta}$ are the parameter vectors to be estimated. The implicit marginal housing price of a forest amenity f_1 is $\partial P_{ij}/\partial f_1 = \delta_1$.

Hedonic wage estimation is conceptually similar. The traditional framework specifies that workers have preferences over a vector of job characteristics (e.g. risk of injury or working conditions). In the case of examining hedonic wages for forest amenities, however, the concern is not with the characteristics of each job, but the landscape characteristics that are tied to the area where one works. The same bundle of forest amenities that may influence hedonic housing prices may compensate workers for lower wages.

Estimation of the hedonic wage must incorporate other factors that affect the wage, namely productivity factors. The most common method is to extend the Mincer (1974) wage equation to include the vector of job characteristics, or in this case, landscape characteristics that includes forest amenities. The hedonic wage equation is,

$$W_{ij} = \pi_0 + \phi' Y_i + \psi' N_j + \zeta' F_j + \mu_i, \tag{12}$$

where W_{ij} is the natural log of individual i's hourly wage, Y_i is a vector of individual human capital characteristics and N_j and F_j are defined above. Neighborhood and forest amenity characteristics should be included to the extent that the geographic areas correspond to labor markets. Note that in the econometric model, i indexes individual wages and characteristics, rather than particular jobs.

The empirical approach adopted in this paper is most similar to Hoehn et al. (1987) and Blomquist et al. (1988), where hedonic regressions for the housing market (equation 11) and wage earners (equation 12) include a common bundle of amenities. The focus of this paper is the importance of forest amenities, so the bundle includes measures of forest area, recreation sites, and other natural amenities that vary across the region.

Hoehn et al. (1987) show that the hedonic results from only one market can be misleading. Depending on the specific assumptions made about the production effects of amenities (i.e. a productive, neutral, or unproductive amenity for firms), the observed sign of the amenity wage and housing price effect can be positive or negative. Thus, any calculation of the implicit amenity price requires both housing and wage differentials.

Equations (11) and (12) can be estimated in linear, logged, log-linear, or some other form.[6] Wage equations, for example, are generally estimated with a log-linear specification (which is derived from its underlying human capital investment theory (Mincer 1974)). In contrast, Hoehn et al. (1987) and Blomquist et al. (1988) use a Box-Cox search over possible specifications to determine the specification for both equations that best fits the data. This latter method is adopted here.

4 Data

The housing price and wage regressions are estimated on a matched sample of wage-earner housing units. That is, each observation has a reported wage, rent or house value, individual characteristics, and housing characteristics. Estimates of equations (11) and (12) are conducted with individual-level data from the 2000 Public Use Microdata Series (PUMS), 5% sample for Arizona and New Mexico (Ruggles et al. 2004). This data gives individual responses to the long form of the 2000 decennial census. Detailed data is available for each respondent on income and earnings, education attainment, race and ethnicity, housing characteristics, labor market participation, and a variety of other variables. Individuals are identified to geographic areas, called Public Use Microdata Areas (PUMAs), with a population of at least 100,000. PUMAs are used to calculate the vector of site-specific characteristics that vary across the landscape and to compile other geographic-based data.

The study region is composed of 51 PUMAs, 36 in Arizona and 15 in New Mexico.

4.1 Dependent variables and population

The two dependent variables are housing price and wages. Housing price is defined as the annual contract rental cost or imputed annual rent. For those households in the sample who rent their dwelling unit, the reported contract monthly rent is multiplied by 12 to obtain annual rent. Households with owner-occupants report the current value of the dwelling. Annual rent is imputed by multiplying this value by an annual discount rate of 7.5%; annual rent for owners is estimated as the income they are earning on that property as an investment, which is assumed to be approximately equal to what they could earn if it was being rented to tenants.[7] Top-coded observations of rent and house value are excluded from the sample (house value is top-coded at $999,998, and rent at $1,700 per month).

The sample for the housing equation is all households from the 2000 PUMS in Arizona and New Mexico who either own or rent their residence (excluding group quarters). Households with imputed observations for ownership status (owner-occupied versus rental), housing value, or rent are excluded.[8] One departure from previous studies is the inclusion of households with 10 or more acres of property. Previous work was generally concerned with urban and suburban residents, so excluding these households ensured that the sample reflected the population and behavior of interest. Although a small portion of our sample (about 1.2%), the population of interest in the intraregional case includes those living in rural areas who may have dwellings on more than 10 acres.

The hourly wage is constructed from reported earnings and work behavior. That is,

$$W_i = wageinc_i/(uhrs_i \times wkswrk_i) \tag{13}$$

where $wageinc_i$ is reported yearly wage income, $uhrs_i$ is usual hours worked per week, and $wkswrk_i$ is number of weeks worked in 1999.

The sample for the wage equation is restricted to wage-earning men and women between the ages of 18 and 64. Those who did not earn wage income, worked fewer than four weeks in 1999, and usually worked fewer than ten hours per week were eliminated from the sample. This sample was created to capture only workers who are closely tied to the wage-earning labor market. Like the housing sample, observations with imputed data were dropped.

The final sample used to estimate (11) and (12) is the intersection of the housing and wage samples. That is, only those observations with complete data under the criteria in both samples are retained for estimation; all observations are household heads under the Census definition. This results in a sample of 36,375 male and 15,285 female household heads.

4.2 Independent variables

The primary interest in the empirical estimates are measures of natural characteristics, and specifically those measures that relate to forest resources. A comprehensive vector of site-specific characteristics has been gathered, including different measures of forest area, recreation sites, water features, hazardous waste sites, recreation-related businesses, and community characteristics. However, several of these measures may exhibit collinearity when included in empirical estimates (see section 5.1). Table 1 describes the geographically-coded independent variables retained for estimation.

The independent variables that measure forest characteristics in a PUMA are the proportion of PUMA land in U.S. Forest Service (USFS) area (AREA_FS) and Congressionally-designated wilderness area (AREA_WILD). USFS area is generally expected to be an amenity, but its designation for multiple uses (including recreation and extractive uses) distinguishes USFS area from wilderness area. Wilderness area, like USFS area, is expected to carry a positive implicit price reflective of recreation, ecosystem services, and passive use values (Phillips 2004). Although these measures are generally thought to be amenable, wildfire risk may offset the positive implicit prices to some extent (Kim and Wells 2005).

Recreation access is measured by the number of recreation sites managed by the USFS (FS_REC) and those managed by other federal agencies (FED_REC). The sum of these two measures (FSFED_REC) is used to estimate shadow prices for recreation sites in general. Recreation sites include camping, hiking, boat launches, and picnic sites.

The other site-specific data used in estimation include square miles of surface water area (SURFACE), number of CERCLA hazardous waste sites (HAZ_COUNT), population density and its square (DENSITY and DENS2), and urban status (URBAN). The forest area measures and surface water area are calculated for each PUMA using GIS software. Hazardous waste site counts and recreation sites are available at the county level; PUMAs are mostly defined as one or more counties, so the county level data is aggregated across the appropriate counties. Recreation site data is gathered from the National Outdoor Recreation Supply Information System (NORSIS), which lists the number of several different types of recreation sites and opportunities for every county in the U.S.[9]

There are three counties that are somewhat problematic for dealing with county-level data: Maricopa and Pima counties in Arizona, and Bernalillo county in New Mexico. These three urban centers are separated into several PUMAs (22 for Maricopa, seven for Pima, and five for Bernalillo county). Given data availability, the approach adopted here is to assign the county-level value to all of the PUMAs in that county.

The independent variables for the housing price equation include categorical variables for property acreage (ACRE9, ACRE10), condominium status (CONDO), number of bedrooms (BEDS), and structure age (AGE). The wage equation independent variables include education (EDUC), potential experience and its square (EXP, EXP2), marital status (MARRIED), immigration status (IMMIG), race indicators (BLACK, NATIVE, ASIAN, RACE2), English language proficiency (NOENG), and transit time to work (TRANTIME). Table 2 describes these variables.

5 Empirical Results

5.1 Estimation issues

Any hedonic estimation involves several econometric issues that must be addressed in order to obtain efficient and reliable results. As mentioned above, a flexible Box-Cox search is used to determine the most efficient functional form. Using maximum likelihood estimation, the generic Box-Cox search estimates,

$$\frac{Y^{\lambda_1} - 1}{\lambda_1} = \alpha + \sum \beta_k \frac{X_k^{\lambda_2} - 1}{\lambda_2} + \epsilon \tag{14}$$

where Y is a dependent variable, X is a vector of k independent variables, and λ_1, λ_2, α, and β are parameters to be estimated. Estimates of $\lambda_1 = 0$ and $\lambda_2 = 0$ correspond with a log-log specification, while $\lambda_1 = 1$ and $\lambda_2 = 1$ is the linear model. Any functional form where the Box-Cox parameters are between -1 and 1 is possible.

The transformation parameter for the continuous independent variables was estimated as $\lambda_2 = 1$. For the housing price equation, λ_1 was estimated between 0.25 and 0.38.[10] The wage equation estimates of λ_1 ranged between 0.007 and 0.05, but the logged form of the wage variable (i.e. $\lambda_1 = 0$) is used to avoid additional computation costs and to allow for simple comparisons of the human capital model to other Mincer-type wage equations in the literature.

There also exists the possibility that wages affect housing prices and housing prices affect wages. As in Blomquist et al. (1988), a fully-simultaneous model is not estimated because calculations of implicit prices using estimations from both markets require reduced form estimates, where the effect of wages on housing prices (and vice versa) is implicit. However, the possibility remains that there are unobserved factors that affect the error terms in both equations. Estimates using a seemingly unrelated regression (SUR) revealed significant, if small, correlation between the error variances of the wage and housing price equations (see Griffiths et al. 1993, 550); for this reason, SUR estimates are used throughout.[11]

Given the large vector of site-specific measures collected as independent variables, another concern is that the data will exhibit multicollinearity. In the case of imperfect multicollinearity, estimates of coefficient standard errors will be large and coefficients sensitive to empirical specification changes (Gujarati 1978, 178). Preliminary regressions using the full vector and sub-vectors of the geographic independent variables exhibited these problems and correlation coefficients between the independent variables confirmed this suspicion. To reduce the consequences of multicollinearity, a subset of geographic variables is sought that are jointly independent.[12] The independent geographic variables described in the previous section, AREA_FS or AREA_WILD, SURFACE, HAZ_COUNT, and FSFED_REC, satisfy the condition of joint independence.

Finally, any econometric exercise that deals with geographic data must consider the possibility of spatially-dependent relationships. Spatial econometric problems arise when data is autocorrelated in space (Anselin 1988). As is the case with time-series data that is autocorrelated, estimates that do not correct for autocorrelation will be inefficient.

The spatial nature of the amenity data collected by PUMA or county suggests that spatial relationships could be present in the PUMS data. Several methods are available to correct for this potential problem. One option is to specify a spatially-autocorrelated model, where error correlation between observations is not zero (i.e., the off-diagonals of the error covariance matrix are not zero). In this case, where the data set yields tens of thousands of observations, such an approach is not feasible because of the size of the spatial weights matrix required for estimation.[13]

The second method is to control explicitly for those spatial characteristics that generate autocorrelation. This approach is more feasible with the current dataset because it only requires additional vectors of independent variables. For example, it may be that the characteristics of near-by places affect housing prices and wages in a particular area; Albuquerque, NM has very few forest amenities within the city limits, but large tracts of wilderness and National Forest areas just outside the city. As a first attempt to control for amenities in close places, the average measure of forest area (AVG_FS or AVG_WILD) in contiguous PUMAs is added as an independent variable for observations in a given PUMA.

5.2 SUR results

The seemingly unrelated regression (SUR) in this context works well in explaining housing prices and wages in the region. Equations (11) and (12) are estimated and reported separately for men and women and reported in tables 3 and 4, respectively.[14] Two primary specifications that are representative of the best fit and have been purged of multicollinearity are reported: one using USFS area measures (AREA_FS and AVG_FS) and the other using wilderness area measures (AREA_WILD and AVG_WILD).

The structural variables in the housing and wage equations are all of the expected sign and highly significant. For example, property acreage and number of bedrooms increases housing value, while years of education and experience increase wages. A test for error correlation between the equations showed that the SUR estimation is indeed justified; a correlation of about 0.20 for men and 0.18 for women was present for all specifications and was always significant.[15]

For both men and women, the proportion of the PUMA in forest area is positive and significantly associated with housing values, and a larger proportion of forest area in contiguous PUMAs is associated with higher housing prices. Further, it appears that forest area in contiguous PUMAs is more important than forests in the household's own PUMA; the housing equation coefficients for AVG_FS and AVG_WILD are larger than AREA_FS and AREA_WILD for men, and AVG_WILD is larger than AREA_WILD for women.

Own-PUMA forest area does not appear to directly affect wages, but nearby forest area is associated with lower wages (a negative coefficient indicates that it is amenable as larger values of that characteristic compensate for lower wages). For both men and women, the coefficients for AVG_FS and AVG_WILD are negative and significant while AREA_FS and AREA_WILD are positive and not significant. In other words, people do not trade wages for forest area where they work, but they will accept lower wages to live and work in an area where forests are closely accessible.

The coefficients for the forest area measures unambiguously indicate that forest area is amenable. But the estimates for the other geographic features are more difficult to interpret. Hazardous waste sites (HAZ_COUNT) are negatively associated with housing prices, which is the expected sign, but the sign is also negative in the wage equation. The coefficients for surface water area (SURFACE) and recreation sites (FSFED_REC) are also the same sign in both equations. Proper interpretation and any meaningful perspective of the magnitude of the results in a SUR framework requires calculations of the total implicit prices.

5.3 Total implicit prices

The above discussion highlights the necessity of considering both the housing and labor markets in tandem. For example, how do we interpret coefficient estimates that are of the expected sign in one market, but the unexpected sign in the other market? What we wish to calculate is the total marginal implicit price for a characteristic, taking into account that characteristic's effect in both markets.

Recall that equation (3) in the theoretical exposition described the implicit demand for amenities and what individuals are willing to trade on the margin for access to amenities. Substituting wages for income in the full model yields

$$\frac{V_f}{V_w} = -\frac{V_p}{V_w}\frac{dp}{df} - \frac{dw}{df}. \tag{15}$$

By Roy's identity and the assumption that the quantity of land purchased for housing is 1, $-\frac{V_p}{V_w} = 1$. Noting that $\frac{V_f}{V_w}$ is the marginal rate of substitution between f and the numeraire good, the implicit price of forest characteristic f is

$$P_f = \frac{dp}{df} - \frac{dw}{df}. \tag{16}$$

The two terms on the right side of equation (16) are simply partial derivatives of the hedonic price functions for housing and labor. Thus, in the case of an amenity, the housing price partial and wage partial need not be positive and negative, respectively, for a total implicit price to be positive.

Calculating implicit prices for the forest amenities begins to clarify some of the econometric results; total annualized marginal implicit prices for the two specifications are reported in the last two columns of tables 3 and 4. As expected from the coefficient estimates, proportion of USFS forest area has a positive marginal implicit price of about \$36 per square mile for men and \$27 per square mile for women. When measured by wilderness area, the implicit prices are higher, between \$70 and \$85 per square mile. Nearby forest area appears to be more "expensive" on the margin, with USFS area in contiguous PUMAs priced between \$80 and \$100 per square mile, and between \$360 and \$400 for wilderness area in contiguous PUMAs. It should be noted that the implicit price for own-PUMA forest area is entirely generated by compensating differentials in the housing market, while contiguous-PUMA forest area prices are generated in both the housing and labor markets.

To put these numbers in perspective, suppose a household moves within the greater Phoenix area, from the north Valley to the west (AZ PUMA 103 to 101). In the specification with the USFS measure of forest area, the move results in an increase in forest area in contiguous areas of about 2.8 percentage points and an annual additional payment of about \$270. A move from Sandoval County, NM to Taos, NM (NM PUMA 500 to 200) increases own-PUMA forest area by 8.5 percentage points and decreases contiguous-PUMA forest area by 1.4 percentage points, resulting in a net payment of about \$1,290 per year.

Although these implicit prices generally support the idea of forest areas as amenities, two questions remain. First, should nearby forest area carry a larger price than forests in one's own PUMA? Most of the households in the sample work in one of the three urban areas of greater Phoenix, Tucson, and Albuquerque. These areas represent the largest source of employment opportunities. But each of these areas is surrounded by nearby forest areas; households may be organizing residential location choices and commuting behavior to live close to amenable forest areas and urban employment opportunities. Under this scenario, it is plausible that nearby forests are more highly valued than forests in one's own PUMA.

On the other hand, it could be that people live in amenable areas and commute to higher wage areas that are disamenable. This would reduce the observed wage price for nearby forest area as households access both high amenities and high wages. Additional regressions using a sample of households that did not commute between PUMAs showed that the AVG_FS and AVG_WILD wage coefficients did not vary from the full sample reported here.[16] This finding is consistent with recent work related to the proximity of forests. Schmidt and Courant (2006) found that the proximity of forests (and other "nice places") near to urban areas is associated with lower wages; people don't need forests in their backyard to accept lower wages, as long as forests are "close enough."

The second question is whether the wilderness area measure should exhibit a larger implicit price than the USFS measure. Recall that the variables selected for the empirical specifications were chosen in part to purge multicollinearity from the data set. To some degree the forest measures are picking up amenable characteristics of other collinear independent variables not directly estimated. Since there is less wilderness area than USFS area (about 2% of the region's land area is in wilderness, while more than 10% is USFS land), smaller changes in the independent variable are picking up the total amenity value of forests as compared to the USFS measure. Thus, it is not unexpected that the wilderness area yields a larger implicit price. This indicates care must be taken when comparing the magnitudes of the implicit prices as each measure may partially proxy for other types of highly-correlated forest characteristics.

Interpretation of the other geographic variables is aided by the calculation of implicit prices, but also raise difficult questions. The implicit price for surface water area is always positive for men and women, between about \$5 and \$15 per square mile per year. This implicit price is paid through lower wages that outweigh lower housing prices in PUMAs with more surface water area. The implicit prices appear to vary by gender for hazardous waste sites and recreation sites. Hazardous waste sites have a negative implicit price for women, but a positive price for men. The opposite is the case for recreation sites; a positive price is observed for men and a negative price for women.

Several plausible explanations for these findings exist that do not contradict the theory and general results in this paper. For example, hazardous waste sites are associated with economic development; while they are probably disamenable, people will seek out areas with good employment opportunities. This phenomenon was demonstrated by Gawande et al. (2000), where net in-migration was shown to be positively associated with hazardous waste sites up to a threshold level where the relationship becomes negative. Thus, a "wrong" sign on the HAZ_COUNT implicit price does not necessarily violate the underlying model.

There are also reasons to believe that men and women will exhibit different attitudes toward environmental characteristics, particularly hazardous waste sites. Bord and O'Connor (1997) observe that women tend to exhibit more concern, on average, than men for environmental issues, and that this gender difference may be due to differential perceptions of risk. In this case, a greater concern for the risks posed by hazardous waste sites by women could account for the negative implicit price estimated for women (between -$12 and -$18) and the positive price for men (up to $18).

Similar reasoning could account for the opposite signs on the implicit prices for recreation sites for men and women. If recreation demand behavior differs significantly by gender, then the implicit prices for recreation sites would also vary by gender. The results reported here are consistent with this reasoning if the gender differences in recreation demand are such that men preferred the type of recreation measured by the FSFED_REC measure.

6 Conclusions

This paper has proposed the open empirical question of whether forest and other natural amenities are important determinants of hedonic housing prices and wages at the regional level. While it has been shown in previous work that forest amenities matter at a local scale, and other characteristics (e.g. climate) matter interregionally, it is not known if regional variation in broad indices of environmental characteristics (e.g. forests) affects housing prices and wages. Results suggest that people make regional housing and labor market decisions based in part on the availability of forest resources, as well as other environmental measures. The proportion of land reserved in public forests, as measured by United States Forest Service land or Congressionally-designated wilderness areas, carries a positive implicit marginal price within the Southwest, meaning that people are willing to pay through the housing and labor markets to live in areas with larger forest tracts. Although no attempt is made here to estimate the welfare impacts of forest areas, it is clear that the annual payment in the region for forest areas is large and significant.

A significant finding is that forests impact both the housing and labor markets, indicating that forests are deeply ingrained in the economy of the Southwest. However, the pattern of this impact differs in each market. Forest area in a household's own geographic area is associated with higher housing prices, while forest area in contiguous geographic areas are associated with higher housing prices and lower wages. People appear to make tradeoffs between several characteristics of real value: access to employment opportunities, commuting time to work, amenable landscapes where they live, and access to nearby amenities.

These results may have important implications for non-market valuation techniques conducted within a region. For example, two of Alan Randall's critiques of the travel cost method (TCM) of valuation are that "recreational preferences may have influenced the choice of residential location," and "the cost of travel time remains an empirical mystery" (Randall 1994, 90). The results presented here suggest that both of these critiques may be important if one seeks to value the characteristics that are generating regional compensating differentials. Individuals appear to be choosing their residential location based on forests, which may be influenced by recreation preferences. And the opportunity cost of travel time (i.e. labor market earnings) is endogenous to the very values a TCM application may seek to elicit. The impact of these results on regional TCM applications deserves further attention.

In general, the results suggest that policies that affect the overall supply of forests can have significant welfare impacts, and the maintenance of forest areas are an important store of value for the quality of life in the region. Less clear is the importance of sub-regional policies that affect the characteristics of public forest lands. Multicollinearity in the data is a barrier to making firm conclusions about the characteristics of specific types of forests or management by particular public agencies. As evidenced by the different implicit prices obtained for hazardous waste sites and recreation sites for men and women, preference heterogeneity may also be an important story within the region.

This leaves open the question of what makes forests amenable. Is it open spaces, recreational opportunities not quantifiable by the number of sites, or the fact that they are publicly-provided lands that matter to people (e.g., culturally, socially, or historically)? Are forests valued for their many uses, or the fact that they provide a beautiful view? Forests provide a vast array of non-market goods and services. Any of these could account for the hedonic premiums that forest areas exhibit, and should be the focus of future research.

A limitation of the findings is related to the spatial nature of the public data used in estimation. Individuals are identified to geographic areas of at least 100,000 people; some of these areas within the study region are very small, while others are large. This raises two potential problems. First, the larger the area is geographically, the more difficult it is to reliably connect individuals within that area to the measured landscape characteristics. Second, the large differences in rural and urban settlement patterns in the region create a difficulty in interpreting the empirical results. That is, the importance of forest areas in small urban areas with small amounts of forest area may be different than in large rural areas with large forested tracts.

Data currently available may not be up to the task of fully capturing the natural amenity story in housing and labor markets at the regional level. The U.S. Forest Service, Southwestern region is currently administering a survey that will address some of these inadequacies by gathering individual survey responses that locate households to particular zip codes or points on the map. These responses will allow for more refined measures of natural amenity access, such as distance to forest sites and the type of land cover surrounding each household. The detailed location information will also allow for direct tests of geographic aggregation bias.

It should also be noted that the estimation of implicit prices occurs in an equilibrium framework. As mentioned at the outset, forest characteristics may be affecting migration patterns within the region. Migration is fundamentally a disequilibrium concept, whereas hedonic theory assumes that utility differences between locations have already been eliminated. Examining the relationship between forest characteristics and migration behavior would help form a more complete picture of the role of these characteristics in the region.

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Table 1: Descriptions of Geographic Independent Variables

Variable	Description (source)	Mean	Std. Dev.
AREA_FS	Percent PUMA area in USFS National Forests and Grasslands (GIS, USFS)	.098	.172
AVG_FS	Mean percent USFS area in contiguous PUMAs (GIS, USFS)	.117	.082
AREA_WILD	Percent PUMA area in Congressionally-designated wilderness areas (GIS, USFS)	.022	.054
AVG_WILD	Area of USFS-defined land in the wildland-urban interface (GIS, USFS)	.024	.023
SURFACE	Area of surface water bodies (excludes streams), sq. miles (Natl. Atlas)	16.02	46.72
HAZ_COUNT	Number of EPA-designated CERCLA hazardous waste sites (EPA)	6.45	7.74
FS_REC	Number of USFS recreation sites (NORSIS)	17.0	9.12
FED_REC	Number of non-USFS federal recreation sites (NORSIS)	20.17	36.47
FSFED_REC	Number of USFS and other federal rec. sites (NORSIS)	37.16	40.08
URBAN	Urban residence: Maricopa AZ, Pima AZ, Bernalillo NM counties (binary, GIS)	.667	.471
DENSITY	Persons per sq. mile in resident's PUMA	1,658.0	1,969.1
DENS2	Square of density	6,625,957	10,500,000

Sources: GIS indicates calculations of polygon- or line-format GIS files for each PUMA.
USFS: United States Forest Service Southwestern Region, http://www.fs.fed.us/r3/gis/datasets.shtml.
National Atlas: http://nationalatlas.gov/atlasftp.html#hydrogm.
EPA: U.S. Environmental Protection Agency CERCLA database, http://www.epa.gov/enviro/html/cerclis/index.html.
NORSIS: National Outdoor Recreation Supply Information System, USFS Southern Research Station.

Table 2: Descriptions of Housing and Wage Structural Variables

		Men		Women	
Variable	Description	Mean	Std. Dev.	Mean	Std. Dev.
Housing variables					
ANN_RENT	Annualized contract rent or dwelling value, $	9,544.4	7,330.0	7,554.3	5,117.3
ACRE9	Property size, 1–9 acres	.122	.328	.087	.282
ACRE10	Property size, greater than 9 acres	.013	.114	.010	.098
CONDO	Condominium status, binary	.016	.126	.044	.205
BEDS	Number of bedrooms	2.76	1.08	2.43	.991
AGE5	Structure age, 2–5 years	.163	.369	.131	.338
AGE10	Structure age, 6–10 years	.111	.315	.098	.297
AGE20	Structure age, 11–20 years	.229	.420	.249	.433
AGE30	Structure age, 21–30 years	.221	.415	.229	.420
AGE40	Structure age, 31–40 years	.096	.295	.101	.301
AGE50	Structure age, 41–50 years	.076	.264	.084	.278
AGE60	Structure age, 51–60 years	.026	.160	.033	.179
AGE61	Structure age, 61 or more years	.026	.158	.034	.182
Wage variables					
LNWAGE	Natural log of hourly wages	2.76	.685	2.50	.646
EDUC	Years of education	13.9	2.82	14.1	2.40
EXP	Years of potential experience, =age-EDUC-6	21.3	11.0	20.1	11.5
EXP2	Square of EXP	575.9	508.6	537.0	505.5
MARRIED	Marital status, binary	.729	.445	.218	.413
IMMIG	Immigrant status, binary	.143	.351	.092	.289
BLACK	Race: Black, binary	.020	.138	.030	.171
ASIAN	Race: Asian/Pacific islander, binary	.018	.133	.014	.117
NATIVE	Race: Native American/Alaska Native, binary	.037	.189	.066	.248
RACE2	Race: Two or more races, binary	.138	.345	.123	.329
NOENG	Speaks English not well or not at all, binary	.046	.210	.024	.152
TRANTIME	Transit time to work, minutes	22.2	19.4	18.8	17.1

Notes: ANN_RENT and LNWAGE are the dependent variables. ANN_RENT is the un-transformed value; results are reported with the appropriate Box-Cox transform.

Table 3: Men, SUR regression results and implicit prices, *obs. = 36,375*

Variable	*Housing* ($\lambda_1 = .25$)		*Wage* ($\lambda_2 = 0$)		*Implicit price ($)*	
ACRE9	2.86***	2.66***				
ACRE10	2.08***	4.16***				
CONDO	-1.03***	-.912***				
BEDS	2.80***	2.81***				
URBAN	2.75***	2.38***				
DENSITY	.001***	.001***				
DENS2	-1.5e-7***	-1.3e-7***				
EDUC			.068***	.068***		
EXP			.033***	.033***		
EXP2			-.0005***	-.0005***		
MARRIED			.145***	.145***		
IMMIG			-.041***	-.041***		
URBAN			.125***	.129***		
BLACK			-.120***	-.120***		
ASIAN			.031	.031		
NATIVE			.071***	-.072***		
RACE2			-.041***	-.040***		
NOENG			-.099***	-.098***		
TRANTIME			.002***	.002***		
DENSITY			2.7e-5***	2.4e-5***		
DENS2			-9.1e-9***	-8.7e-9***		
AREA_FS	5.32***		.039		36.30	
AVG_FS	7.17***		-.084*		80.86	
AREA_WILD		12.6***		.045		85.73
AVG_WILD		32.8***		-.373**		365.01
SURFACE	-.005***	-.005***	-.0004***	-.0005***	14.78	16.51
HAZ_COUNT	-.167***	-.149***	-.003***	-.003***	3.76	12.78
FSFED_REC	.006***	.006***	4.9e-5	8.5e-5	4.59	5.00
R^2	.3888	.3862	.2188	.2189		
Residual corr.					.1985	.1991

Notes: Structure age variables are suppressed for brevity. ***, **, and * indicate 99%, 95%, and 90% confidence levels, respectively. Residual correlations between the equations are significant at the 99% level. Implicit prices are calculated as follows: *Housing:* $\frac{\partial p_L}{\partial f} = b_{f_h} * ann\_rent^{(1-\lambda_1)}$ where b_{f_h} is the coefficient for characteristic f from the housing equation and ann_rent is the calculated annual rent evaluated at the sample mean. *Wages:* $\frac{\partial w}{\partial f} = b_{f_w} * wage\_inc$ where $wage\_inc$ is annual wage earnings evaluated at the sample mean. The prices for the area variables are converted to per square mile figures using average PUMA area. All implicit prices are annualized.

Table 4: Women, SUR regression results and implicit prices, *obs. = 15,285*

Variable	*Housing* ($\lambda_1 = .38$)		*Wage* ($\lambda_2 = 0$)		*Implicit price ($)*	
ACRE9	5.32***	5.22***				
ACRE10	10.2***	10.4***				
CONDO	-3.06***	-2.65***				
BEDS	5.44***	5.42***				
URBAN	4.81***	4.14***				
DENSITY	.003***	.002***				
DENS2	-3.8e-7***	-3.0e-7***				
EDUC			.086***	.086***		
EXP			.030***	.030***		
EXP2			-.0005***	-.0005***		
MARRIED			.014	.013		
IMMIG			8.6e-5	.0002		
URBAN			.154***	.168***		
BLACK			-.068**	-.069***		
ASIAN			.056	.057		
NATIVE			.032	.030		
RACE2			-.060***	-.058***		
NOENG			-.058*	-.055*		
TRANTIME			.002***	.002***		
DENSITY			7.8e-6	6.9e-6		
DENS2			-3.3e-9*	-3.0e-9*		
AREA_FS	12.8***		.022		27.52	
AVG_FS	11.6***		-.301***		97.98	
AREA_WILD		33.2***		.010		71.46
AVG_WILD		50.1***		-1.19***		396.58
SURFACE	-.015***	-.016***	-.0002***	-.0003***	4.16	5.84
HAZ_COUNT	-.536***	-.501***	-.004***	-.004***	-18.55	-11.67
FSFED_REC	.007**	.010***	.0004***	.0004***	-9.25	-10.37
R^2	.2946	.2910	.2191	.2196		
Residual corr.					.1839	.1844

Notes: Structure age variables are suppressed for brevity. ***, **, and * indicate 99%, 95%, and 90% confidence levels, respectively. Residual correlations between the equations were significant at the 99% level. Implicit prices are calculated as follows: *Housing:* $\frac{\partial p_h}{\partial f} = b_{f_h} * ann\_rent^{(1-\lambda_1)}$ where b_{f_h} is the coefficient for characteristic f from the housing equation and *ann_rent* is the calculated annual rent evaluated at the sample mean. *Wages:* $\frac{\partial w}{\partial f} = b_{f_w} * wage\_inc$ where *wage_inc* is annual wage earnings evaluated at the sample mean. The prices for the area variables are converted to per square mile figures using average PUMA area. All implicit prices are annualized.

Notes

[1]See Niemi et al. (1999) for a discussion of the second paycheck. Also, Garber-Yonts (2004) provides a thorough review of literature related to the impact of natural characteristics on migration, hedonic prices, and local economic conditions.

[2]Hand (2006) investigated labor market compensating differentials for linguistic enclave amenities in the Southwest, but did not incorporate regional variation in natural amenities.

[3]The regional planning process was set forth in the final rule for National Forest System Land Management Planning, Federal Register, v.70 n.3, January 5, 2005, pg. 1055.

[4]See "Southwestern Region Revision Plan Strategy" April 21, 2006. http://www.fs.fed.us/r3/plan-revision/strategy.shtml, accessed 9/19/2006.

[5]While shown as a scalar, the term f could also represent a vector.

[6]See Freeman (2003, ch.11) for a discussion of different functional forms for housing models.

[7]The 7.5% figure was used in Costa and Kahn (2003). Hoehn et al. (1987) and Blomquist et al. (1988) use a discount rate of 7.85% estimated from a separate source.

[8]The PUMS data often uses reported characteristics to impute other missing data. These are dropped since the characteristics used to impute data are often the same ones used in the hedonic regressions.

[9]The NORSIS data is maintained by Carter Betz at the USFS Southern Research Station. Contact information available at http://www.srs.fs.usda.gov/trends/betz.html.

[10]Several sub-samples of data are used in estimation, and the estimate for λ_1 varies between these samples. The appropriate value of λ_1 is displayed in the results tables.

[11]The results reported in (Blomquist et al. 1988) do not use a SUR model, although the authors note that the cross-equation error correlations are significant. The coefficient estimates

reported here were more sensitive to the restriction of error independence than in Blomquist et al. (1988).

[12]The method for selecting variables follows the suggestion of Gujarati (1978, 184), where ordinary least squares regressions of one independent variable on the other variables are used to identify significant linear relationships between the independent variables. A set of variables is acceptable if the regression of each variable on the other variables all do not reject joint independence. All F-tests for the combinations in the results did not reject independence at the 95% confidence level.

[13]Kim et al. (2003) estimated spatial hedonic models with spatial data organized similarly to the data used here. The authors specified a contiguity-based spatial weights matrix, where all observations within a sub-district (like a PUMA in our data) were considered neighbors. This same approach is conceptually possible with the data in this paper, but the 609 total observations in Kim et al. (2003) are significantly more manageable than the 50,000-plus observations used here. A promising related method that deserves attention in this context has been developed by Conley (1999).

[14]Interacting a categorical variable for gender with all of the other independent variables confirmed that both the wage and housing price equation should be estimated separately for men and women (i.e., all of the gender interactions were jointly significant). See Griffiths et al. (1993, 419) for the relevant F-statistic.

[15]The Breusch-Pagan test statistic is used to determine the significance of correlation (Griffiths et al. 1993).

[16]Results available upon request.

**Stated Preferences for Ecotourism Alternatives
On the Standing Rock Sioux Indian Reservation**

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Abstract

Despite favorable locations and the potential for economic development, Native American tribes have not developed their ecotourism markets substantially. This paper presents a choice experiments analysis of potential tourist and local resident preferences for alternative ecotourism development scenarios for the Standing Rock Sioux Indian Reservation. The choice experiments' elicitation featured attributes of both cultural and nature-based tourist attractions. Survey results demonstrated that visitors interviewed at powwows had significantly different preferences from those interviewed at local tourist attractions. Results from all samples showed positive preferences towards an amphitheater, a nature trail, and a bison meal, and no preference toward an ATV trail. Non-powwow tourists had significant willingness to pay for a number of potential attractions, including nature trails, a road through the bison pasture, and an interpretive center with amphitheatre show.

Keywords: choice experiments, ecotourism, Native Americans, Standing Rock Sioux

Studies have shown that ecotourism is the fastest growing segment of the international tourism market (Lew, 1996). Despite the potential for economic development and being located in areas that are rich with natural beauty and near other tourist destinations, Native American tribes have not developed their ecotourism markets adequately to capitalize on this increasing market demand. Only a few reservations have made efforts to diversify tourist opportunities, beyond gaming, and broaden visitorship (Lew, 1996). Correspondingly, there has been little published research on the demand for ecotourism Native American reservations.

Ecotourism, also known as nature-based tourism, is defined as "tourism that consists of traveling to relatively undisturbed or uncontaminated natural areas with the specific objective of studying, admiring, and enjoying the scenery and its wild plants and animals, as well as any existing cultural manifestation found in these areas (Ceballos-Lascurain, 1987, in Fennel 2001)". Ecotourists can be thought of as tourists who demonstrate stewardship to cultures and to the environment. As a result, ecotourism can offer an economic return to the host communities for conserving and celebrating their cultures. Ecotourism development promises to offer indigenous peoples employment alternatives which complement the natural beauty of reservation lands and respects Native American cultural traditions (Wearing and Neil, 1999).

This paper presents a case study of an analysis of the preferences among potential tourists and local residents for alternative ecotourism development scenarios for the Standing Rock Sioux Indian Reservation (SRSIR). Choice experiments are used to assess preferences and estimate the willingness to pay of tourists for hypothetical ecotourism packages. Both the nature-based and culture-based attractions are assessed.

Thus, this study provides an opportunity to assess not only the willingness to pay of potential tourists for ecotourism services, but the preferences of SRSIR residents towards the tourist services they prefer to be offered. It also provides a means to compare interest in natural and cultural attractions. The paper thus contributes in a number of ways to the literature on ecotourism.

The second part of this paper provides background on the SRSIR. This is followed by a short literature review on the economic analysis of ecotourism. The subsequent sections of this paper provide details of the methodology employed and the results of the analysis. The paper concludes with recommendations for SRSIR tourism authorities.

Background

The SRSIR is the home of the Lakota band of Sioux Indians. The reservation was established in 1889 in the wake of the Great Plains Wars (Tiller, 1996). It encompasses all of Sioux County, North Dakota and Corson County, South Dakota and is governed by the Standing Rock Sioux Tribal Government. According to the 2000 US Census the reservation has a population of 8,241, with a median family income of $23,922. Forty percent of the population remains below the poverty level. The total land area of the Standing Rock Sioux Indian Reservation is 2.3 million acres, and of that, 1,408,061 million is tribally owned (The Confederation of American Indians, 1986). The land is primarily occupied by short grass prairie. Buttes, some with elevations of up to 2,000 feet, are common throughout the lands (Tiller, 1996).

The SRSIR has a number of tourist amenities, including Lake Ohae, the Cannonball River, Fort Manual Lisa, Fort Yates, and Sitting Bull's original and reestablished graves. Lodging and meals are available at the reservation's two casinos as well as a number of smaller facilities. Highway 1806, which traverses the SRSIR, is a gateway to the Teton Sioux Nation and crosses four Sioux Indian Reservations. It links cultural and recreational sites throughout North and South Dakota and was named a Native American Scenic Byway in 2005. It has many historical sites and monuments (see Figure 1). Standing Rock Sioux Tribal Tourism (SRSTT) promotes visits to the reservation. Tours which feature historical background and visits to the buffalo pasture are offered to groups and individuals. A number of Native American artists are promoted by the SRSTT and periodic art fairs are held. The Tribe and its districts host a number of powwows which are social gatherings and cultural events which include social and ceremonial dances, traditional costumes, and competitions. These powwows are open to the public and promoted to tourists. Hunting and fishing is welcome, with landowners' permission and the appropriate Tribal license.

Despite the promotion of tourist visits to its powwows and attractions, SRSTT admits that some tribal members might be uncomfortable with increased tourism. The SRSTT brochure on visitor etiquette stresses many common courtesies, such as requests to not trespass on private land nor litter. Additional requests include to demonstrate respect for elders and to refrain from direct eye contact and photography during ceremonies. The brochure cautions tourists to respect sacred sites including unmarked graves, and to refrain from attending certain ceremonies unless invited.

Literature Review

There is some scholarly research on tourism on tribal reservations. Lew (1996) used a survey of tribal authorities within the United States to assess the administrative practices dedicated to tourism and tourism promotion. Lew concludes that ecotourism development on tribal reservations is not as successful as it could be. With rapid growth in international cultural tourism during the 1990s, the author advocated that tribes need to restructure their tourist industry initiatives to capitalize on this trend. Schneider and Salk (2004) administered onsite questionnaires to assess visitor interest in cultural and nature-based experiences on Leech Lake Band of Ojibwe Reservation. The authors concluded that the potential experiences that attracted the highest interested respondents are traditional Native American dance performances, tribal gift shops, and Native American cultural heritage history centers. Browne (1989) used published and survey data to assess the economic development from reservation tourism and concluded that the economic motive for developing or maintaining a reservation tourism industry remains strong. In many cases, tourism development seems to be related to increased self-esteem, self determination, in addition to positive economic growth (Browne, 1989).

Research on ecotourism in North Dakota is limited. Hodur et al (2004) and Leistritz et al (2004) assessed opportunities for ecotourism development in North Dakota and southwest North Dakota respectively and concluded that outdoor recreation opportunities that featured hunting, fishing water sports, nature watching and birding had the most growth potential.

Research on ecotourism has generally stressed its potential in promoting the preservation of natural, cultural, and historical places (Luzar *et al.,* 1995). Mieczkowski (1995) and Boo (1990) provide overviews that highlight both financial and environmental benefits. Some empirical studies have highlighted the positive impacts of ecotourism. Wunder (2000) showed that tourism increased local income and provided incentive to support conservation in Ecuador. Lindberg (1996) assessed ecotourism at a number of protected areas in Belize and concluded that tourism generated net financial benefits for local residents and support for conservation. However without additional user fees it did not generate positive net financial support for protected area management.

A growing body of literature has used stated preference techniques to assess willingness to pay for different ecotourism experiences. Kelly *et al.* (2006) used a discrete choice experiment (CE) method to examine visitor preferences for land-use, transportation, recreation, and other environmental initiatives intended to promote eco-efficiency in tourism destinations. Hearne and Salinas (2002) assessed preferences of local and international tourists for ecotourism development options in Costa Rica. Lindberg *et al.,* (1999) used choice experiments to assess residents' attitudes towards the costs and benefits of increased tourism on a community. Hearne and Santos (2005) assessed tourists' and local residents' preferences towards protected area management strategies in Guatemala.

Methodology

Choice experiments are a stated preference technique that allows analysts to assess preferences and estimate willingness to pay from respondents' responses to a hypothetical market solicitation. Choice experiments are based upon two theoretical foundations, Lancasterian consumer theory and random utility theory. Lancasterian theory posits that utility is derived from the attributes of a particular product. Random utility theory posits that individual utility ($\boldsymbol{U}$) is unknown but can be decomposed into a systematic or deterministic component ($\boldsymbol{V}$) and an unobserved or stochastic component ($\boldsymbol{\varepsilon}$). Thus, for individual $\boldsymbol{j}$ in scenario $\boldsymbol{i}$, utility can then be expressed as

$$\boldsymbol{U_{ij} = V_{ij} + \varepsilon_{ij}}. \tag{1}$$

Since the systematic component can be expressed as a linear function of explanatory variables, $\boldsymbol{V_{ij}}$, can be referred to as

$$\boldsymbol{V_{ij} = \beta' \mathrm{x}_{ij}}. \tag{2}$$

The analysis of multiattribute choice experiment data requires maximum likelihood estimation. Assuming independently and identically distributed Type 1 extreme value error terms with a scale factor μ and a variance σ^2, where $\mu > 0$ and $\sigma^2 = \pi^2/ 6\mu^2$, it is possible to use the multinomial logit model, such that the conditional probability of alternative $\boldsymbol{A}$ being selected out of a set of alternatives $\Phi = (\boldsymbol{A}, \boldsymbol{B}, \boldsymbol{C})$ is estimated as

$$P(A|\Phi) = \frac{\exp(\mu V_A)}{\sum_j \exp(\mu\, V_j)} \qquad \forall j \in \Phi\,. \tag{3}$$

The multinomial logit model requires the assumption of independence of irrelevant alternatives (IIA), which implies that the probability of choosing one alternative over another is unaffected by the presence or absence of additional alternatives (Louvierre et al, 2000; Hensher et al, 2005).

The nested multinomial logit model is used when the scenarios are logically grouped into a decision tree and the respondents' decision making process is seen to be iterative. In this case, a respondent must first decide whether to opt for an ecotourism visit package or for *no visit*. If an ecotourism package is chosen, then the respondent can decide which of the presented ecotourism packages to select. One advantage of the nested logit model is that it does not require the IIA assumption. The nested logit model assumes that an individual's probability of choosing a new proposed alternative $\boldsymbol{i}$ is a function of the probability of choosing any new alternative, as opposed to the *no visit* option, as well as the preference toward alternative $\boldsymbol{i}$ over the other proposed alternatives in the choice set $\boldsymbol{J_s}$. Thus, the proposed trip alternatives are considered to be nested into one branch, $\boldsymbol{s}$, in a decision tree that includes an alternative branch, $\boldsymbol{n}$, for *no visit* (see Figure 2). Assuming an extreme value distribution of the error term in the utility function, this probability can be expressed as:

$$P_{is} = P(i|s)P(s) = \left[\frac{\exp(V_{is}|\alpha_s)}{\exp(I_s)}\right]\left[\frac{\exp(\alpha_s I_s)}{\sum_{k=s,n}\exp(\alpha_k I_k)}\right] \text{ with} \tag{4}$$

$$I_s = \log\left[\sum_{i=1}^{J_s}\exp(V_{is}/\alpha_s)\right]. \tag{5}$$

where ***P(s)*** is the probability of choosing a new scenario, ***P(i |s)*** is the probability of choosing alternative ***i*** once the decision to choose a new scenario was made, V_{is} is the indirect utility of alternative ***i***, α_s is the inclusive value coefficient which measures the substitutability across alternative tourist products. I_s is known as the inclusive value and is a measure of the expected maximum utility of the alternatives J_s (Green, 2003; Kling and Thomson, 1996).

As an initial phase of the research, an experts' meeting was held to provide an understanding of research needs and local concerns, to identify attributes for analysis in the choice experiments, and to identify survey procedures. Local experts stressed that there has always been a certain niche demand for cultural tours of the SRSIR. These experts also suggested that the reservation's natural attractions could be used to diversify and lengthen tourists' visits. They also stated that many tribal members may be apprehensive towards increased tourism.

Later a series of focus groups was held with tribal members, tourists, and entrepreneurs. Focus group protocol, as established by Krueger (1988), was followed throughout the focus group process. Focus group meetings were held with: audience members at the Kenel, South Dakota powwow; nature-based tourists in Mobridge, South Dakota; tourists at Fort Rice State Historic Site; campers at Sugar Loaf State Park; various residents in a number of the reservation communities; visitors to a tribal art symposium; employees of Sitting Bull College; and employees of the Grand River Casino.

These focus groups identified certain favored attractions, such as an amphitheater, a demonstration farm tour, and an ATV trail. Some individuals stressed the need for family activities. Based upon these meetings, a preliminary survey instrument was developed and conducted among tourists and residents at a local powwow. After the results of the preliminary survey were analyzed, attributes and levels were chosen for empirical analysis. Table 1 presents the attributes and levels that were used in the final survey. Both natural attractions and cultural attractions were selected. The prices used correspond to the per person price of a tour package which includes the attributes of the choice profile. The price levels of $80.00 to $200.00 are within the range of $55 per hour per person charged by Standing Rock Tribal Tourism for guided historical tour (Standing Rock Tribal Tourism, undated).

The full factorial experimental design, of $4^4 * 2^3$ combinations, was reduced with an algorithm that maximizes D efficiency to produce 432 choice profiles (Zwerina et al., 1996). The combinations of attributes forming each scenario, and the combination of choice scenarios forming each choice set, were chosen for their fulfillment of the following criteria: (1) orthogonality, which aims at ensuring that the attributes vary independently one from another between scenarios; (2) level balance between attributes, meaning that the different levels of each attribute appear with equal frequency among the choice scenarios; and (3) minimal overlap between levels of each attribute within a choice set. The fourth criteria, utility balance between alternatives, could not be taken into account because of the lack of prior information on the public preferences for the different possibilities of PES spreading presented. These criteria are conditions to be used for the estimation of the parameters associated with each attribute when considering an underlying linear utility function. These choice profiles were then grouped into 108 choice profile combinations. Each choice profile combination included three choice profiles, listed as *A*, *B*, and *C*, as well as a fourth option of *No Visit*.

The survey instrument was designed to be brief, in order to minimize the time spent by a respondent to complete it. Respondents were asked a few questions about their interest in tourism on the SRSIR and a number of demographic questions. Each respondent was asked four choice experiment solicitations. An information package was also developed in order to ensure that there was consistent information presented to the respondents. Each attribute and attribute level was explained.

Three separate populations were considered for analysis: SRSIR residents, tourists who visit cultural and natural amenities of the SRSIR, and tourists who visit sites proximate to and similar to the SRSIR. Surveying was conducted by one of the coauthors and a locally recruited enumerator in August and September of 2006. A number of local tourist sites, both on and off the reservation, were selected for surveying. Fort Yates was considered to be a convenient spot for surveying local residents, since it is an administrative area for the whole reservation. In addition a number of powwows were used for surveying because they serve as gathering places for residents and tourists. Table 2 presents the distributions of the sample across various sites. Respondents were approached, given preliminary information on the survey, and asked if they were willing to participate. Participants were handed a clipboard with information on the SRSIR and the survey. These respondents completed the survey in the presence of the enumerator (see Tuscherer, 2007 for details of the surveying procedure).

Ecotourists on the reservation were difficult to encounter, therefore this population was combined with the off-reservation tourist population. However, a number of tourists were encountered at various powwows in the region. These were later considered separately for statistical analysis. Two hundred and five potential respondents were asked to complete the survey. There were one hundred and eighty-three willing respondents. One hundred and forty-two surveys were deemed usable: 54 locals, 54 powwow tourists, and 34 non-powwow tourists. Table 3 presents the residency of the respondents. Data was analyzed using LIMDEP NLogit 3.0 (Greene, 2002).

Results

Multinomial logit models, as presented in Louviere *et al* (2000) and Hearne and Salinas (2002), were estimated for the two samples, residents and tourists. A likelihood ratio test as described by Swait and Louviere (1993) was used to test the difference in preference orderings between powwow and non-powwow tourists. The equality of the combined coefficients and scale parameters was rejected with the following test:

$$-2[LL\text{ (pooled tourist data)} - LL\text{ (powwow)} - LL\text{ (non-powwow)}] \qquad (5)$$

where LL is the log likelihood function, which is distributed χ^2 with 14 degrees of freedom for the number of restricted parameters. The calculated value of $\chi^2_{14} = 22.76$ ($p = 0.064$) is greater than the 21.07 critical value to reject equality with 90% confidence. Following procedures presented in Swait and Louviere (1993) the relative scale factor $\mu_{non\text{-}powwow}/\mu_{powwow}$ was estimated to be 0.90 and the data for the powwow sub-sample was adjusted. The log likelihood test was then rerun with the adjusted data set and the calculated value of $\chi^2_{14} = 21.44$ ($p = 0.091$) is greater than the 21.07 critical value to reject equality with 90% confidence. Thus the preference orderings of the powwow and non-powwow populations are considered to be unequal and are listed separately in the subsequent models.

Table 5 presents results of the three multinomial logit estimations. The coefficients for the alternative specific constants (ASC) for choices *A*, *B*, and *C* show the preference for choosing one of these alternatives over the *No Visit* alternative. Clearly the samples of residents and powwow tourists have positive preferences for any of the hypothetical trip alternatives over *No Trip*. Each of the other variables listed in the model, except *Price*, have been coded as discrete variables. Thus, the coefficients represent a preference over the unnamed 'no' alternatives, such as: *No farm/ranch visit*, *No bison processing*, *No herd visit*, and *No trail*. Results of this model demonstrate that all three populations have positive and significant preferences for a visit which features a bison meal, a combination bison meal and tanning class, a stagecoach ride through a bison pasture, a nature trail, and an interpretive center with an amphitheater show. All populations showed no significant preference towards ATV trails. Residents demonstrated little interest in any of the culinary farm/ranch tour options. But they did have interest in a *Hide Tanning Class*. Non-powwow tourists had little interest in a *Hide Tanning Class.*

A number of nested logit models were tested. All demographic variables were tested for significance within the first level decision of whether or not to accept a hypothetical ecotourism package. Results from the selected nested logit model, with the first level decision of ecotourism participation as a function of age, education, and days dedicated towards tourism are presented in Table 6. These results were used in series of a likelihood ratio tests as described by Louviere et al. (2000) to test if the nested model has better explanatory power than the multinomial logit model. Results of these tests are shown below.

$$2[LL \text{ (nested local)} - LL \text{ (multinomial local)}] = 16.88 \sim \chi^2_7; \quad (6)$$

$$2[LL \text{ (nested powwow)} - LL \text{ (multinomial powwow)}] = -1.42 \sim \chi^2_7; \quad (7)$$

$$2[LL \text{ (nested non-powwow)} - LL \text{ (multinomial non-powwow)}] = 17.05 \sim \chi^2_7. \quad (8)$$

The 7 degrees of freedom are for the added restrictions on the nested model. Given that the calculated value would need to be greater than 12.02 in order to reject the equality of the two models with 90% confidence, the nested model is considered to be superior to the multinomial logit model for the sample of locals and non powwow tourists.

Table 5 shows mostly similar results to the results of the multinomial models. In all three models the alternative specific constants for options *A*, *B*, and *C*, were, as expected, insignificant and are not reported. The first level decision of whether or not to accept a hypothetical trip package is a function of education level, age, and annual tourism days. Among the residents and the non-powwow tourists, higher educated respondents and those that spend more time in tourism were less likely to respond with *No Visit*. Older non-powwow respondents are less likely to choose one of the ecotourism alternatives.

The important difference between the populations is the preference towards lower prices. As expected, the local population did not have a significant preference for lower prices. This is not unexpected because many local respondents would not expect to pay this fee themselves. Instead, they might believe that these prices would be paid by outside tourists and become income into the reservation. Also the powwow tourist did not have a significant preference towards lower prices. This is somewhat surprising, because it does not conform to economic theory. However it does conform to previous literature that suggests that certain cultural tourists have a high willingness to pay for certain activities (Moscado and Pearce, 1999). It is also possible that powwow attendees are internalizing the concerns of the tribal residents who may be providing services as opposed to internalizing the concerns of tourists who would be buying the services. The last group of non-powwow visitors did have a highly significant preference towards lower prices.

Marginal willingness to pay (MWTP) was estimated for only the sample of non-powwow tourists. These were surprisingly high. The statistically significant MWTP estimates included: \$145 for a drive through the bison pasture; \$118 for a nature trail; \$105 for a culinary farm tour with a cooking class; \$102 for a stagecoach ride through the bison pasture; and \$102 for an interpretive center with an amphitheater show. These relatively high MWTP estimates could be due to a relatively small sample size. It is also possible that the one sub-population with a significant preference towards spending less money could be misrepresenting their true WTP because of a warm glow effect, which at the time of the response gives the respondent satisfaction from hypothetically doing the right thing.

Conclusions

The objective of this study was to assess preferences for and willingness to pay for additional ecotourism attractions on the SRSIR. Initial efforts to sample three separate populations were thwarted by the absence of ecotourists visiting the SRSIR.

However analysis of the data demonstrated that among tourists, the sub-population of tourists that were interviewed at powwows had significantly different preference ordering than non-powwow tourists interviewed at local historical and recreation sites.

The key difference among the results for the different samples was the preference towards lower prices. Local residents were indifferent towards prices. This is not surprising given that residents might expect not to pay for ecotourism, but to directly and indirectly benefit from tourist dollars entering the reservation. Powwow tourists had the same indifference towards prices as local residents. Non-powwow tourists significantly preferred lower prices which allows for a reliable estimation of willingness to pay.

Both multinomial logit and nested logit models were estimated. In general, the nested logit models showed more explanatory power. The results showed positive preferences towards increased ecotourism option on the SRSIR. Results from all samples demonstrated positive preferences towards an amphitheater, a nature trail, and a bison meal. Each sample had no preference toward an ATV trail. Tourists favored a road through the bison pasture but locals had no significant preference for this. Locals favored a hide tanning class while the non-powwow tourists did not favor this option. Willingness to pay was estimated for the sample of non-powwow tourists. The estimated values were within the range of the prices currently charged for guided history tours of the SRSIR. These results are in line with Lew's (1996) study which indicated that ecotourism on Indian Reservations is underdeveloped. SRSIR tourism personnel should view ecotourism development as offering alternatives to destructive industries as well as offering new employment opportunities while maintaining the natural beauty of their lands and preserving their Native American cultural traditions.

This research should assure reservation tourism personnel that the local population supports the development of ecotourism alternatives on the reservation. Indeed, this overwhelming support concurs with the Lindberg *et al.* (1999) study which indicated that residents are willing to accept tourism development, with potential negative impacts, provided that they also receive positive impacts. The overall highest respondents' preference is toward an interpretive center with an amphitheater show. This result is consistent with the Schneider and Salk (2004) study that indicated Native American cultural heritage history centers as being among the top three interests of respondents in their study.

Table 1: Attributes and Levels of Choice Sets

ATTRIBUTE	LEVELS
Demonstration farm/ranch	1. Culinary farm/ranch tour 2. Culinary farm/ranch tour and hands-on cooking class 3. Culinary farm/ranch tour and cattle round-up 4. No farm/ranch visit
Bison Processing	1. Hide tanning class 2. Authentic bison meal 3. Authentic bison meal and hide tanning class 4. No bison processing
Bison Herd Visit	1. Driving road through herd pasture 2. Stagecoach ride through herd pasture 3. No herd visit
Trails	1. Nature trail 2. Bike trail 3. ATV trail 4. No trail
Tribal history	1. Interpretive signs at highway pullouts 2. Interpretive center and amphitheater show 3. No history presentation
Price	1. \$80.00 2. \$120.00 3. \$160.00 4. \$200.00

Table 2. Survey Application

LOCATION	SURVEYS COMPLETED	POPULATION
Fort Yates, ND	28	Local residents (all eight districts represented)
Wakpala, SD Powwow	13	Local residents (8) Reservation tourists (5)
Fort Berthold, ND Powwow	20	Non-Reservation ecotourists
Mobridge, SD	6	Local residents (5) Reservation tourists (1)
Grand River Casino Resort, SD	6	Local residents (4) Reservation tourists (2)
Fort Mandan, ND	25	Non-Reservation ecotourists
Knife River Indian Village, ND	6	Non-Reservation ecotourists
Fort Abraham Lincoln, ND	11	Non-Reservation ecotourists
Bismark, ND	16	Local residents (7) Non-Reservation ecotourists (9)
United Tribes Powwow Bismarck, ND	52	Local residents (15) Non-Reservation ecotourists (37)

Table 3: Respondents' location of residence

	Number	Percent
SRSIR	43	30.3
North Dakota	40	28.3
South Dakota	5	3.5
Minnesota	8	5.6
Other US	32	22.5
Europe	1	0.7
Other country	4	2.8
Other tribe	9	6.3

Table 4: Results of Multinomial Logit Models

	SRSIR Residents (n = 216)		Powwow Tourists (n = 216)		Non-Powwow Tourists (n = 136)	
	Coefficient *Standard Error*		**Coefficient** *Standard Error*		**Coefficient** *Standard Error*	
ASC Trip ‘A’	-1.464	***	-2.989	***	-0.963	
	0.528		*0.582*		*0.668*	
ASC Trip ‘B’	-1.068	**	-2.572	***	-1.320	*
	0.518		*0.562*		*0.678*	
ASC Trip ‘C’	-1.128	**	-2.871	***	-1.288	*
	0.525		*0.581*		*0.692*	
Culinary farm/ranch tour	0.213		0.560	**	0.660	*
	0.268		*0.282*		*0.371*	
Tour and cooking class	0.163		0.763	***	0.842	**
	0.256		*0.277*		*0.352*	
Tour and cattle round-up	0.237		0.632	**	0.643	*
	0.259		*0.289*		*0.378*	
Hide tanning class	0.604	**	0.563	*	0.110	
	0.283		*0.273*		*0.364*	
Bison meal	0.957	***	0.649	**	0.720	**
	0.274		*0.278*		*0.343*	
Meal and tanning class	1.108	***	0.819	***	0.685	**
	0.217		*0.281*		*0.355*	
Road through bison pasture	0.217		0.594	**	1.057	***
	0.235		*0.247*		*0.320*	
Stagecoach through bison pasture	0.429	*	0.824	***	0.746	**
	0.224		*0.240*		*0.326*	
Nature trail	0.692	***	0.939	***	0.833	**
	0.264		*0.278*		*0.345*	
Bike trail	0.528	*	0.845	***	0.512	
	0.273		*0.283*		*0.354*	
ATV trail	0.298		0.421		-0.263	
	0.271		*0.290*		*0.372*	
Signs at highway pullouts	0.236		0.997	***	0.128	
	0.231		*0.261*		*0.317*	
Amphitheater show	0.620	***	1.426	***	0.748	**
	0.219		*0.263*		*0.308*	
Price	-0.000		0.000		-0.008	***
	0.002		*0.002*		*0.003*	
Significance of the model $\chi^2(14)$	41.76	***	72.67	***	48.44	***

*,**, *** significant at the 90%, 95%, 99% confidence level (P[|Z|>z])

Table 5: Results of Nested Logit Models

	SRSIR Residents (n = 216)	Powwow Tourists (n = 216)	Non-Powwow Tourists (n = 136)
	Coefficient *Standard Error*	**Coefficient** *Standard Error*	**Coefficient** *Standard Error*
First Level Decision Visit or No Visit			
Education Level	0.448 ***	-0.022	0.476 **
	0.164	*0.163*	*0.200*
Age	-0.073	0.135	-0.966 ***
	0.154	*0.145*	*0.318*
Annual Tourism Days	0.311 *	0.097	0.319 **
	0.165	*0.119*	*0.138*
Second-Level Decision Attributes of Trip			
Culinary farm/ranch tour	0.385	0.487	0.658 *
	0.292	*0.302*	*0.397*
Tour and cooking class	0.290	0.808 ***	0.863 **
	0.269	*0.301*	*0.371*
Tour and cattle round-up	0.345	0.590 *	0.635
	0.275	*0.308*	*0.399*
Hide tanning class	0.741 **	0.551 *	0.085
	0.294	*0.282*	*0.388*
Bison meal	0.932 ***	0.772 **	0.721 **
	0.289	*0.303*	*0.365*
Meal and tanning class	1.108 ***	0.887 ***	0.770 **
	0.281	*0.300*	*0.392*
Road through bison pasture	0.306	0.647 **	1.200 ***
	0.251	*0.270*	*0.369*
Stagecoach through bison pasture	0.517 **	0.865 ***	0.841 **
	0.241	*0.260*	*0.361*
Nature trail	0.715 **	1.001 ***	0.976 **
	0.286	*0.303*	*0.437*
Bike trail	0.520 *	0.895 ***	0.652
	0.296	*0.308*	*0.425*
ATV trail	0.281	0.335	-0.225
	0.287	*0.300*	*0.393*
Signs at highway pullouts	0.284	0.911 ***	0.105
	0.240	*0.273*	*0.339*
Amphitheater show	0.694 ***	1.327 ***	0.837 **
	0.228	*0.279*	*0.357*
Price	0.000	-0.000	-0.008 ***
	0.002	*0.002*	*0.003*
Inclusive Value Parameters			
No Visit	1.000	1.000	1.000
	fixed	*Fixed*	*Fixed*
Visit	-0.022	0.315	0.670
	0.141	*0.247*	*0.419*
Significance of the model $\chi^2(21)$	164.0	221.2	100.1

*,**, *** significant at the 90%, 95%, 99% confidence level (P[|Z|>z])

Table 6: Marginal Willingness to Pay for Non-Powwow Tourists

Significant Attribute Levels

	MWTP	Standard Error	
Farm/ranch tour and cooking class	$105.03	57.8	*
Bison Meal	$87.82	50.9	*
Bison meal and hide tanning class	$93.72	55.2	*
Driving road through herd pasture	$145.79	66.0	**
Stagecoach ride through herd pasture	$102.39	54.6	*
Nature trail	$118.78	68.6	*
Interpretive center and amphitheater show	$101.88	57.8	*

*,** significant at the 90%, 95% confidence level (P[|Z|>z])

Figure 1: The Standing Rock Sioux Indian Reservation

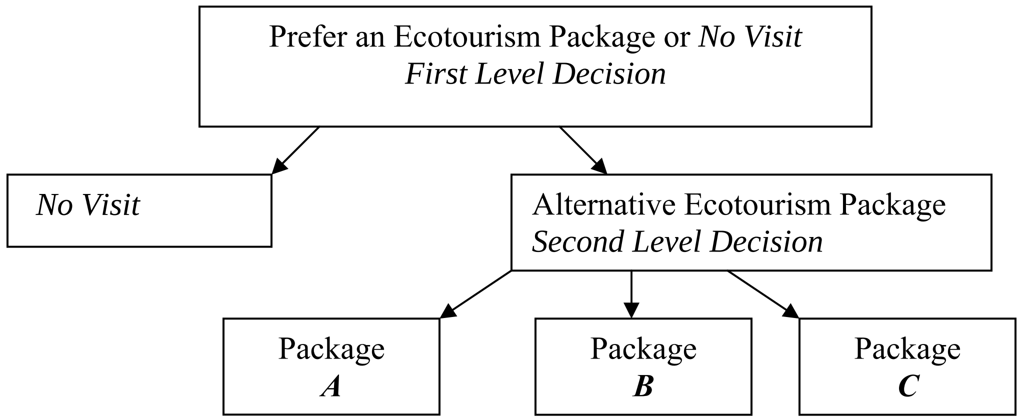

Figure 2: The Nested Decision Making Structure

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Meta-Functional Benefit Transfer for Wetland Valuation: Making the Most of Small Samples

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Meta-Functional Benefit Transfer for Wetland Valuation: Making the Most of Small Samples

Abstract:

This study applies functional Benefit Transfer via Meta-Regression Modeling to derive valuation estimates for wetlands in an actual policy setting of proposed groundwater transfers in Eastern Nevada. We illustrate how Bayesian estimation techniques can be used to overcome small sample problems notoriously present in Meta-functional Benefit Transfer. The highlights of our methodology are (i) The hierarchical modeling of heteroskedasticity, (ii) The ability to incorporate additional information via refined priors, and (ii) The derivation of measures of model performance with the corresponding option of model-averaged Benefit Transfer predictions. Our results indicate that economic losses associated with the disappearance of these wetlands can be substantial and that primary valuation studies are warranted.

Key words: Bayesian Model Averaging, t-Error Regression Model, Meta-Analysis, Benefit Transfer, Wetland Valuation

JEL codes: C11, C15, Q51

I) Introduction

There is an ongoing trend amongst resource managing agencies to use existing information on past policy outcomes to predict economic benefits for planned policy implementations. For some institutions this concept of "Benefit Transfer" (BT) has essentially become the primary tool of policy assessment. For example, in a recent insiders' assessment of the role of BT at the U.S. Environmental Protection Agency (EPA) Iovanna and Griffiths (2006) predict that due to the triple constraints of expediency, financial strains, and administrative hurdles "original assessment studies will remain a rare exception" in future EPA valuation efforts.

It is not surprising, therefore, that the concept and methodology of BT has become of central interest to resource economists in the U.S. and abroad. In recent years much effort has been allocated to examine the theoretical underpinnings of BT (e.g. Smith and Pattanayak 2002; Smith et al. 2006) and to facilitate its econometric and computational implementation (e.g. Rosenberger and Loomis 2000a; Léon et al. 2002; Moeltner et al. 2007; Moeltner and Rosenberger 2007). Most of these studies have illustrated their respective methodologies using simulated data or hypothetical policy scenarios. However, to date there are very few contributions to the published literature that show-case the implementation of BT within the context of an actual policy implication.[1]

The first objective of this study is to provide a "real-world" example of BT within the context of wetland valuation. Existing meta-analyses of wetland valuation studies (Brouwer et al. 1999; Woodward and Wui 2001; Borisova-Kidder 2006; Brander et al. 2006), have focused primarily on the marginal effects of wetland functions and attributes on economic benefits. In contrast, this analysis illustrates how BT *was actually* used to inform decision makers on projected economic benefits related to a planned resource intervention affecting wetlands.

The second aim of this study is to demonstrate how Bayesian econometric techniques can overcome many classical estimation challenges when even the most thorough literature review produces only a small number of existing sources suitable for a meta-functional BT model. Specifically, we illustrate how (i) study-specific heteroskedasticity – ubiquitous in meta-regressions – can be addressed with only a single additional model parameter, (ii) ancillary information from source studies and meta-analyses can be used to derive more informed Bayesian priors, and (iii) Bayesian Model Averaging (BMA) can be employed to address model uncertainty in the derivation of BT predictions. We believe that our approach will widen the applicability of BT in many resource valuation settings, and provide resource planners with a tool to proceed with meta-functional BT in many cases where its implementation has to date been hampered by small sample problems.

We introduce the general policy context that triggered our BT implementation in the next Section, Section III describes the construction of the meta-dataset underlying our analysis, Section IV discussed the econometric framework and Section V presents estimation results. We summarize our findings and offer concluding remarks in Section VI.

II) Policy Context and Study Area

The Southern Nevada Water Authority (SNWA) has recently proposed a Groundwater Development Project to transfer approximately 200,000 acre-feet of groundwater per year from seven hydrographic basins in rural Eastern Nevada to the wider Las Vegas Valley to assure a reliable water supply for this fast growing urban area in future years. The construction of groundwater conveyance facilities and supporting infrastructure is proposed to start as early as 2009 (SNWA 2006).

The granting of associated water rights rests with the Nevada State Engineer ‘s Office. Over the last two years, this agency has collected scientific and economic evidence on the potential implications of this water transfer for the targeted provider areas to aid in this decision-making process. The due date for all evidence to be submitted was November 1, 2006. Approximately three weeks prior to this deadline the authors were contacted to examine if there might be any “non-market” – type economic values that could be at stake should the Project be approved.

Given the tight time frame and available scientific evidence we decided to focus on two distinct and unique wetland areas, the Swamp Cedar Natural Area and the Shoshone Ponds Natural Area. Both wetlands are located in Spring Valley in east-central NV. They are predicted to be significantly and almost immediately impacted by the planned groundwater withdrawals (Lanner 2006).

Spring Valley is approximately 9.5 miles wide (east – west) and 95 miles long (north-south). It distinguishes itself from other valleys in the Great Basin by its high elevation (5500 – 6000 feet), and its relatively abundant water resources, provided by over 100 natural springs (Charlet 2006). These springs together with snowmelt retained by a hardpan soil layer support numerous wetlands throughout the Valley (Lanner 2006).

The Swamp Cedar Natural Area (SCNA), a marshy ecosystem with natural ponds and meadows contains 3200 acres of public land administered by the Bureau of Land Management (BLM). It supports a large stand of Rocky Mountain Junipers (Juniperus scopulorum), commonly referred to as "Swamp Cedars". These Spring Valley Cedars have been described as "globally unique" as they have adapted to a distinctly different environment than is characteristic for the main population of their species (Charlet 2006; Lanner 2006). The SCNA offers recreational opportunities for hiking, primitive camping, and wildlife viewing, although it does not feature a designated access road, parking area, developed trail system or established campgrounds (BLM 1980a).

The Shoshone Pond Natural Area (SPNA) is located approximately 13 miles south of the SCNA in Southern Spring Valley. It contains 1240 acres of public land managed by the BLM. It features two important natural resources: (i) A second stand of "Swamp Cedars" of the same ecotypical variety as those found in the SCNA, and (ii) Three manmade, spring-fed pools that harbor two rare species of fish, the relict dace (Relictus solitarius) and the Pahrump poolfish (Empetrichthys latos). The relict dace is listed by the Nevada Natural Heritage database as "imperiled and vulnerable in Nevada and globally", while the Pahrump poolfish, for which the Shoshone ponds constitute one of only three remaining habitats, has been federally listed as an endangered species since 1969. While lacking maintained hiking trails or established campsites, the SPNA offers recreational opportunities for hiking, primitive camping, and wildlife viewing (BLM 1980b).

Both time frame and available funds were insufficient to launch a primary valuation study, which left BT as only viable alternative to produce at least approximate estimates of potential economic losses. Furthermore, given that some information on basic attributes was available for these wetlands we aimed for BT via function transfer (e.g. Kirchhoff et al. 1997; Brouwer and Spaninks 1999). In addition, since there does not exist a single valuation study that corresponds sufficiently well to the current context we decided to implement this functional BT via a Meta-Regression Model (MRM) that draws information from several underlying source studies (e.g. Rosenberger and Loomis 2000b; Shrestha and Loomis 2001; Johnston et al. 2005).

III) Data Set Construction

Suitable primary studies for the MRM were identified using the following sources: Four existing meta-analyses focusing on the economic value of wetlands (Brouwer et al. 1999; Woodward and Wui 2001; Brander et al. 2006; Borisova-Kidder 2006), the Environmental Valuation Reference Inventory (EVRI), a searchable database focusing on non-market valuation, and ECONLIT, a general searchable database for economic literature. The initial criteria for study selection were: (i) Geographic area = USA or Canada, (ii) Exclusion of coastal or marine types of wetlands, (iii) Estimated economic values must include values related to habitat, biodiversity, or species preservation. The latter two criteria flow from the nature of the current policy context: Spring Valley wetlands are distinctly different ecosystems than coastal or marine wetlands, and their economic value is primarily related to habitat and biodiversity services. Thus, we excluded studies that focused on wetlands with the *sole functions* of flood control or water quality improvements, as well as studies that *only* examined the value of specific wetlands with respect to extractive use (hunting, fishing).

This "first cut" approach produced a set of 24 initial candidate studies. Given the nature of their primary valuation objectives (habitat and biodiversity services, recreational opportunities) all of these sources use survey-based approaches to elicit households' willingness-to-pay (WTP) to preserve or expand a specific existing wetland area. A second round of screening eliminated studies that are based on identical wetlands and target populations, and studies based on surveys that produced response rates below 30 percent. In the case of duplicate studies we retained the study with the most reliable research methodology. The low-response rate criterion was applied to guard against "selection bias", i.e. the possibility that the small segment of those who participated in the survey is not representative of the underlying target population. Only one study fell into that latter category.

These selection refinements resulted in a final set of nine studies deemed suitable for the research context at hand, yielding 12 observations available for our meta-dataset (One study, Blomquist and Whitehead 1998, reports WTP estimates for four different wetlands). While this sample is not as large as would be ideal it has several desirable properties. As shown in the Table 1 the selected studies provide good coverage of the geographic target area, with applications from various parts of the United States, and one Canadian contribution (Tkac 2002). All studies were conducted within the last 15 years and thus use modern survey and estimation methodologies. The underlying target populations are of a general nature with at least regional scope. Specifically, three studies (Loomis et al. 1991; Roberts and Leitch 1997; Tkac 2002) focus on a regional population of stakeholders, while five of the studies are associated with a State-wide target population (Hanemann et al. 1991; Whitehead and Blomquist 1991; Mullarkey 1997; Blomquist and Whitehead 1998; Poor 1999) and one source (Klocek 2004) has nation-wide coverage.

The sample also exhibits a desirable mix of journal publications, book chapters, government reports, and theses or dissertations. The relatively strong representation by contributions from the "gray" literature eases the traditional concern of "publication bias" in meta-analytical research, i.e. the notion that only valuation results that are surprising or otherwise noteworthy are ever considered by journal editors.

Table 2 provides more detailed information for each observation included in our meta-dataset. Most policy scenarios presented to respondents for a given study stipulated that wetland areas would be lost (due to agricultural activities, mining, or urban sprawl) if no action was taken. With two exceptions (Roberts and Leitch 1997, who use a payment brackets-approach, and Tkac 2002 who uses a payment table with uncertainty scales á la Welsh and Poe 1998), all of the studies employed a variant of the dichotomous choice elicitation format. In most cases respondents were then asked if they would be willing to pay a specific dollar amount ("bid") into a nature conservation fund or in additional taxes to preserve these lands. The only exceptions to this “loss avoidance” approach are the studies by Poor (1999) and Klocek (2004) who asked respondents if they would be willing to contribute to a special fund to *create* additional acres of wetland (for example by converting drained agricultural areas to their original marshy conditions). The relative homogeneity of elicitation approaches is fortuitous for our meta-application since the small sample size would preempt the inclusion of study-methodological regressors, which have been found to carry considerable importance in MRMs (e.g. Johnston et al. 2005, Moeltner et al. 2007) . Given that most of our source studies use a similar elicitation approach we can argue that the effect of methodological aspects should be minor for our MRM.

All of the included studies asked respondents to value the entire bundle of wetland services, including habitat and biodiversity provision, flood control, water filtration, and opportunities for non-consumptive (wildlife viewing, hiking, photography) and consumptive (hunting, fishing) recreational activities. Some studies (Blomquist and Whitehead 1998, Tkac 2002) also stress the presence of threatened or endangered species on the wetlands under consideration. Since the surveys targeted the general population of underlying households (as opposed to a specific group of active users), only a relatively small segment of respondents indicated that they had visited the wetland under consideration in the past, as depicted in the "percentage of active users" column in Table 2. Thus, the lion's share of estimated economic benefits (i.e. reported WTP) is likely associated with non-use or existence values. This is another important and desirable feature of our data set given the current research context, since it can be expected that only a small proportion of the wider population of stakeholders will have actually visited the Spring Valley wetlands considered in this study.

As evident from the table the types of wetland, the policy scenarios in terms of wetland acres preserved or created, and the percentage of active users in the underlying sample vary widely over studies. Not surprisingly, so does aggregate WTP per household. The smallest welfare estimate (in 2006 currency) is less than a dollar per U.S.-wide household to preserve parts of the Canaan Valley National

Wildlife Refuge in West Virginia, while households in the San Joaquin Valley in California are willing to pay close to $300 per year to prevent the loss of a large share of the Valley's wetland habitats. In consequence, using any single study for a point transfer of benefits would be risky. Furthermore, using the sample mean of over $60 would likely lead to grossly inflated BT predictions given the relatively obscure and isolated nature of our policy site. In combination, these facts support the approach of functional BT via meat-regression.

IV) Econometric Framework

Classical Challenges

Virtually all existing MRMs on resource valuation have been estimated via classical least squares methods. However, in our case the small number of observations makes it difficult to take a classical estimation approach. For example, any specification test relying on asymptotic theory will be unreliable in this case. This includes all tests on heteroskedasticity, a likely occurrence in meta-regressions given that each source observation flows from a different original regression model. On the other hand, simply ignoring heteroskedasticity would cast doubt on the reliability of standard errors for estimated coefficients and associated confidence intervals for BT predictions. Traditionally, researchers have addressed heteroskedasticty problems with robust, or White-corrected, standard errors (e.g. Woodward and Wui 2001; Brander et al. 2006). However, the White correction itself rests on asymptotic theory and is thus of limited value in our application. Similarly, unless the analyst firmly believes that the basic Classical Linear Regression Model applies, there are no specification tests available for guidance on the composition or functional form of explanatory variables.

Related problems arise when the MRM is estimated with logged WTP as dependent variable, as has been the case in virtually all existing meta-analyses related to resource valuation to assure non-negativity of welfare measures, but absolute values in dollars are required for BT predictions. The 'Delta method' (e.g. Greene 2003, Ch.5) or equivalent asymptotic techniques such as bootstrap or the popular

Krinsky and Robb (1986) routine must be applied to obtain confidence intervals for the converted estimates. Again, such approaches are unreliable in our case.

We thus propose a Bayesian estimation framework for our MRM as it poses several advantages over a classical approach in this context. First, error heteroskedasticity can be modeled hierarchically with only a single additional parameter. Second, relevant information form source studies that is not captured in the actual meta-data can enter the model via the specification of prior distributions, which leads to more representative and, possibly, more efficient BT estimates. Third, a Bayesian framework allows for the incorporation of *model uncertainty* by estimating multiple candidate models and their corresponding probability weights, and then deriving a weighted model-averaged distribution for the benefit construct of interest. This circumvents the need to choose, with little guidance, a single preferred specification to generate the transfer function, as would be required in a classical framework.

Bayesian Model

Our kernel MRM takes the form of a standard linear regression model with the added feature of a hierarchical distribution for the variance of the regression error to allow for observation-specific heteroskedasticity. Specifically, we stipulate individual variance weights to be drawn from an inverse-gamma distribution with equal shape and scale. Our baseline MRM can thus be written as

$$y_j = \mathbf{x}_j'\boldsymbol{\beta} + \varepsilon_j \quad \text{with} \quad \varepsilon_j \sim n\left(0, \sigma^2 \omega_j\right), \quad \text{and} \quad \omega_j \sim ig\left(\tfrac{v}{2}, \tfrac{v}{2}\right), \tag{1}$$

where y_j is WTP reported in study j, $\mathbf{x}_j$ is a vector of population and wetland characteristics associated with study j, $\boldsymbol{\beta}$ is a corresponding vector of regression coefficients, ε_j is a zero-mean regression error with variance $\sigma^2\omega_j$, and ig denotes the inverse-gamma distribution.[2]

As shown in Geweke (1993) the hierarchical specification of the variance of ε_j is exactly equivalent to drawing ε_j from a t-distribution with mean zero, scale σ^2 and v degrees of freedom. In addition to capturing variance-inequalities across observations, this allows for higher probabilities of large error variances than would be expected for a basic normal model, a likely occurrence in a meta-regression context. To be specific, for any given σ^2 a small value of v (say 5 to 10) implies a heavy-tailed distribution, while, as is well known, the *t*-distribution approaches normality for larger values of *v*.

Conditional on the individual-specific variance weights the likelihood function for our MRM follows a multivariate normal distribution with generalized variance-covariance matrix, i.e.

$$pr\left(\mathbf{y} \mid \mathbf{X}, \boldsymbol{\beta}, \sigma^2, \boldsymbol{\omega}\right) = \left(2\pi\right)^{-n/2} \left|\left(\sigma^2 \boldsymbol{\Omega}\right)\right|^{-1/2} \exp\left(-\tfrac{1}{2}\left(\mathbf{y} - \mathbf{X}\boldsymbol{\beta}\right)' \left(\sigma^2 \boldsymbol{\Omega}\right)^{-1} \left(\mathbf{y} - \mathbf{X}\boldsymbol{\beta}\right)\right)$$
$$\text{with} \quad \mathbf{X} = \left[\mathbf{x}_1' \quad \mathbf{x}_2' \quad \cdots \quad \mathbf{x}_n'\right]', \quad \boldsymbol{\omega} = \left[\omega_1 \quad \omega_2 \quad \cdots \quad \omega_n\right], \text{ and } \quad \boldsymbol{\Omega} = diag\left[\omega_1 \quad \omega_2 \quad \cdots \quad \omega_n\right] \tag{2}$$

where n denotes the total number of observations. Conditioning the likelihood on $\boldsymbol{\omega}$ instead of v facilitates posterior simulation. In essence, we treat the variance weights as additional data. This is deemed 'data augmentation' in Bayesian analysis (Tanner and Wong 1987).

The specification of the Bayesian model is completed by assigning prior distributions to all model parameters. We follow standard approaches by choosing a multivariate normal distribution with mean $\boldsymbol{\mu}_0$ and variance-covariance matrix $\mathbf{V}_0$ for the vector of regression coefficients $\boldsymbol{\beta}$, and an inverse-gamma distribution with shape η_0 and scale κ_0 for the shared variance component σ^2. In addition, we specify the heteroskedasticity parameter v to follow a gamma distribution with shape 1 and inverse scale $1/v_0$. As discussed in Koop (2004), Ch. 6, this choice of hyper-prior distribution for v is computationally convenient and assures the required condition of $v > 0$. Thus, the hierarchical prior structure for our MRM can be compactly denoted as

$$\begin{aligned}
&pr\left(\boldsymbol{\beta}\right) = mvn\left(\boldsymbol{\mu}_0, \mathbf{V}_0\right) \\
&pr\left(\sigma^2\right) = ig\left(\eta_0, \kappa_0\right) \\
&pr\left(\omega_j \mid v\right) = ig\left(\tfrac{v}{2}, \tfrac{v}{2}\right), \forall j \qquad p\left(v\right) = g\left(1, 1/v_0\right)
\end{aligned} \tag{3}$$

The Bayesian framework then combines likelihood function and priors to derive marginal posterior distributions for all parameters. We use a Gibbs Sampler (GS) with a built-in Metropolis-

Hastings (Hastings 1970) routine for draws of *v* along the lines suggested in Koop 2004, Ch. 6, to simulate these distributions. The details of this algorithm are given in Appendix A.

Model weights and BT predictions

As described in more detail in the next section we estimate a set of *M* different MRMs, distinguished by the specification of explanatory variables and chosen parameters for prior distributions. For each model M_m, $m = 1 \ldots M$, the GS generates $r=1 \ldots R$ draws of parameters $\boldsymbol{\beta}$, σ^2, and v. For notational convenience we combine these parameters into a joint vector $\boldsymbol{\theta}$ and denote individual draws of this vector as $\boldsymbol{\theta}_\mathbf{r}$. For each model, our ultimate construct of interest is the posterior predictive distribution (PPD) of WTP for the policy context, conditional *only* on policy site regressors and the general model specification, i.e. $pr\left(y_p \mid \mathbf{x}_\mathbf{p}, M_m\right)$, where subscript *p* indicates "policy context". The derivation of this density proceeds in two steps: First, for each round of the original GS we obtain a draw of y_p conditional on a specific set of parameters, ie. a draw from $pr\left(y_p \mid \mathbf{x}_\mathbf{p}, \boldsymbol{\theta}_\mathbf{r}, M_m\right)$, as $y_{p,r} = \mathbf{x}_\mathbf{p}'\boldsymbol{\beta}_\mathbf{r} + \varepsilon_{p,r}$ where the error term $\varepsilon_{p,r}$ is drawn from $t\left(0, \sigma_r^2, v_r\right)$. Second, we repeat this process *S* times to obtain multiple draws of y_p for each set of original parameter draws.[3] The resulting $\left(R \cdot S\right)$ draws can thus be interpreted as flowing from $pr\left(y_p \mid \mathbf{x}_\mathbf{p}, M_m\right)$, the desired simulated PPD of y_p for model M_m.

For each model the posterior simulator also produces the *model-conditioned marginal likelihood* i.e. $pr\left(\mathbf{y} \mid M_m\right)$, which can be used to compute the *posterior probability* for a given model, denoted as $pr\left(M_m \mid \mathbf{y}\right)$. Loosely speaking this probability indicates how likely the observed data (i.e. the dependent observations in our MRM) were generated by model M_m. Formally, the two concepts are linked through Bayes' Rule as

$$pr\left(M_m \mid \mathbf{y}\right)=\frac{pr\left(\mathbf{y} \mid M_m\right) pr\left(M_m\right)}{p\left(\mathbf{y}\right)} \tag{4}$$

where $pr\left(M_m\right)$ indicates the *prior model probability*, and $p\left(\mathbf{y}\right)$ denotes the unconditional marginal likelihood, i.e. the probability that **y** was generated by *any* of the considered models. Assuming that the M models considered for our application constitute an exhaustive representation of all possible models (i.e. $\sum_{m=1}^{M} pr\left(M_m \mid \mathbf{y}\right)=1$), and setting $pr\left(M_m\right)=1/M, \forall m$, we can write the posterior model probability as the ratio of a specific model's marginal likelihood to the sum of all marginal likelihoods for the considered model space, i.e. as

$$pr\left(M_m \mid \mathbf{y}\right)=\frac{pr\left(\mathbf{y} \mid M_m\right)}{\sum_{m=1}^{M} pr\left(\mathbf{y} \mid M_m\right)}, \tag{5}$$

(e.g. Koop 2004, Ch. 1). Since an analytical expression for $pr\left(\mathbf{y} \mid M_m\right)$ does not exist for our kernel specification we simulate this value for each model using the approach proposed by Chib and Jeliazkov 2001.

Model-specific PPDs of the BT construct and probability weights can then be combined to produce a model-averaged predictive distribution of y_p. Analytically, this density can be expressed as

$$pr\left(y_p \mid \mathbf{x_p}\right)=\sum_{m=1}^{M}\left\{\int_{\boldsymbol{\theta}} pr\left(y_p \mid \mathbf{x_p}, M_m, \boldsymbol{\theta}\right) pr\left(\boldsymbol{\theta} \mid \mathbf{y}, \mathbf{X}, M_m\right) d\left(\boldsymbol{\theta}\right)\right\} pr\left(M_m \mid \mathbf{y}\right). \tag{6}$$

Equation (6) indicates that the posterior predictive distribution of y_p conditional only on policy descriptors $\mathbf{x_p}$ is derived by marginalizing conditional draws of y_p over (i) model parameters, and (ii) all considered models. The first marginalization is accomplished via the two-step approach outlined above. Marginalizing over model space simply involves weight-averaging model-specific draws from the PPD of y_p across all models.[4]

V) Empirical Implementation

Model space and prior refinements

Table 3 captures the model space, i.e. the exhaustive list of all MRMs considered in our analysis. We decided on this set based on plausible and available explanatory variables, available extraneous information for prior distributions, and preliminary estimation runs.[5]

The models in Table 3 differ in three dimensions: (i) The distribution of the error term, (ii) The specification of prior distributions for $\boldsymbol{\beta}$, and (iii) The functional specification of the explanatory variable "wetland acreage". As indicated in the table, Models 1- 4 have a t-distributed error as given in equation (1), while the error term in Models 5-8 follows a generic homoskedastic normal distribution. All models indexed by "a" or "b" specify a prior distribution for the constant term with a mean of zero and a variance of 10. Based on preliminary OLS results, we also allow for specifications with a large negative mean (-50) and a highly diffuse variance for the constant term (all models indexed by "c" in the table). Prior distributions for the slope coefficients contained in $\boldsymbol{\beta}$ are non-informative for "a"-type models, and refined for "b" and "c"-type MRMs. Wetland acreage enters the models in one of four possible forms: logged and logged-squared (Models 1 and 5), logged only (Models 2 and 6), linear in units of 1000 plus linear-squared (Models 3 and 7), and linear only (Models 4 and 8).

The remaining regressors include annual household income, in log-form, and the percentage of active wetland users corresponding to a given source study. Information for the refinement of priors was available for income and wetland acreage. This added information flows primarily from coefficient estimates reported in source studies or other meta-analyses. The detailed process of prior refinement is described in Appendix B. As shown in the appendix, our refined priors have considerably smaller variances than their diffuse, or non-informative, counterparts. Thus, they carry substantial weight in the posterior simulation routine and allow the added information to have a measurable effect on posterior results.

Estimation results

All models are estimated using 15000 burn-in draws and 10000 retained draws in the Gibbs Sampler. The decision on the appropriate amount of burn-ins was guided by Geweke's (1992) convergence diagnostic (CD). For each MRM, the standard deviation of the proposal density for v in the Metropolis Hastings algorithm contained in the GS (denoted as s_v in Appendix A) is set to achieve an optimal acceptance rate of 44-50% (see e.g Gelman et al. 2004 Ch. 11).

Table 4 presents estimation results for each model in terms of its logged marginal likelihood (denoted as $log\ pr(y|M)$), mean absolute percent error (MAPE), and model weight (indicated as $pr(M|y)$ in the table)[6]. As discussed in Kass and Raftery (1995) marginal likelihood values are primarily useful to compare two competing models (the ratio of marginal likelihoods is often referred to as "*Bayes Factor*") or to generate model weights for each specification in a set of ex ante chosen candidates, which is the main focus in our application. The MAPE, in turn, is a universal measure of model fit that focuses primarily on the in-sample predictive ability of a given specification. For a general discussion of this and other measures of model fit in a Bayesian framework see Gelman et al. (2004), Ch. 6. Moeltner et al. (2007) compare different measures of model fit in an empirical setting.

According to Kass and Raftery (1995), a difference in marginal likelihoods of 5 or more indicates a decisive superiority of the model with the smaller value (in absolute terms). Accordingly, as can be seen from Table 4, our models with homoskedastic-normal errors score considerably better using this criterion than their t-error counterparts. This likely indicates that our data set is too small to provide substantial evidence of error heteroskedasticity. Applying equation (5), higher marginal likelihood values also translate into higher posterior model weights for the normal specifications as indicate in the last column of the table. In fact, none of the t-error models receives any appreciable posterior weight in our application.

Within each group of error distributions models with refined priors and zero-mean prior for the constant term (i.e. "b"-type models) receive slightly more favorable marginal likelihood scores than their respective non-informative versions ("a"-type models) or versions with a large negative intercept prior ("c"-type models). This effect is especially pronounced for Models 4b vs. 4a and Models 8b vs. 8a, with likelihood differences in the 4-5 point range. Not surprisingly, the refined priors generally lead to a deterioration in predictive ability with respect to the actual data as indicated by the higher MAPE scores for "b" and "c" type models compared to their "a" type counterparts. As mentioned above, our refined priors with their smaller variances absorb considerable posterior weight, while the diffuse priors for "a" –type models allocate most posterior weight to the underlying data. In general, this leads to better in-sample predictive ability, i.e. lower (= better) MAPE scores. The exceptions to this pattern can be again observed for Models 4b and 8b and, to a lesser extent, Models 4c and 8c, which receive a better MAPE score than their corresponding non-informative specifications. It thus appears that the linear formulation of wetland acreage *without* its squared form produces the best model fit for our refined MRMs. Overall, Model 8b receives by far the highest posterior model probability (0.902) and would thus be an ideal candidate to generate BT predictions if a single model had to be chosen for this task.

However, we pursue our original strategy of allowing every model to contribute to the posterior distribution of BT estimates by computing the weighted average of model-specific results. We generate predicted benefits for three possible groups of stakeholders: (i) All households in the four counties surrounding Spring Valley (11,118 units by the 2000 Census), (ii) All Nevada households (751,165 units), and (iii) All households in Nevada and Utah (1,452,446 units). We set income figures to the most recently reported Census medians, and the percentage of active users arbitrarily to 5% for the Four County-Region and 1% for wider Nevada and Utah. Naturally, a small poll of households from the target population could produce more accurate estimates of active use. Our chosen figures are at the lower end of those found in our source studies and thus appear reasonably conservative for the task at hand.

We follow the steps outlined in Section IV to generate predictive distributions. For each of the R = 10,000 parameter draws from the original GS, we draw a set of S = 100 predicted values for policy outcome y_p. We then keep every 20^{th} of these draws to reduce autocorrelation in our sequence. Thus, we retain 50,000 posterior predictive draws for our analysis.[7]

Table 5 captures posterior predictive distributions of benefits for the Four Counties ("Region"), Nevada ("NV") and Nevada plus Utah ("NV/UT") produced by each single MRM. Model-averaged results are given toward the bottom of the table. For each model the table reports the posterior mean and its numerical standard error (nse), a measure of simulation noise.[8] As illustrated e.g. in Koop (2004), Ch. 3, a numerical 95% confidence interval for the posterior mean can be obtained as [posterior mean $\pm 1.96 \cdot nse$]. As can be seen from the Table posterior means range from under \$3/ year (Models 3a, 7a) to over \$30 (Model 1c). Thus, in absence of any additional information on relative model performance it would be risky to base BT predictions on a single MRM. This clearly illustrates the benefits of guidance through posterior model weights or the convenience of model-averaging.

Using our model-averaged results we thus predict annual losses associated with the disappearance of the SCNA and SPNA wetlands of \$4.8 to \$5.6 per household. The numerical confidence intervals for these posterior means are in the \$0.1 - \$0.2 range, indicating that simulation error is a minor consideration for our application. These per-household figures translate into predictions of \$62,000 for the Region, \$3.6 million for Nevada, and close to \$7 million for Nevada and Utah as indicated in the bottom row of the table.[9]

VI) Conclusion

This study describes an actual application of meta-functional BT to value wetland areas in Eastern Nevada. We illustrate how Bayesian estimation techniques can be used to produce reasonable BT results even with a very small underlying sample of source studies. The main advantages of our methodology compared to classical regression models are (i) The ability to capture heteroskedasticity in straightforward fashion through hierarchical modeling of the error variance, (ii) The ability to incorporate additional information not captured in the data via refined priors, and (ii) The availability of measures of model performance with the corresponding option of generating model-averaged BT predictions.

While the hierarchical error variance approach turned out to be of limited importance in our application, the refinement of prior distributions led to clearly superior posterior results for several of our specifications. In addition, guidance through marginal likelihood values and associated posterior model weights proved critical in identifying promising MRMs and in properly weighting individual models in the generation of model-averaged predictive distributions.

Naturally, we also ought to stress the limitations of our approach compared to a primary valuation study for the policy area. Our small sample size and the lack of detailed information on specific attributes of wetland areas considered in original studies preempts a more thorough examination of the effect of various wetland features (other than acreage) on WTP. Each of the wetlands in our meta-data is unique in some sense, and wetland size in acres alone is not necessarily a reliable proxy for wetland quality attributes. For example, it is quite possible that the Spring Valley wetlands are valued more highly than predicted in our analysis given their function as habitats for a globally unique stand of trees, and two threatened / endangered fish species. On the other hand, many of the included wetlands in our meta-regression offer richer recreational opportunities than the Spring Valley areas. This, in turn, could inflate our BT estimates.

Perhaps the most meaningful way to interpret our secondary-data results is to use them as a strong indication that the economic losses associated with a potential disappearance of Spring Valley wetlands could be of substantial magnitude, and that therefore primary economic research is both warranted and justified. Given the large geographic scale of the proposed groundwater extraction project, and the potentially irreversible nature of its environmental implications, it is imperative that decision makers be informed of all economic benefits and costs involved. These considerations should also include non-market type values associated with affected natural areas. We hope that our preliminary results via BT will aid in creating awareness that such values exist and that they can be of important magnitude.

Appendix A: Posterior Simulation

This Appendix outlines the detailed steps of the Gibbs Sampler (GS) for the regression model with t-distributed errors. It is convenient to apply Tanner and Wong 1987's concept of data augmentation and treat draws of $\boldsymbol{\omega} = \left[\omega_1 \quad \omega_2 \quad \cdots \quad \omega_n\right]$ as additional data in the likelihood function. This leads to the augmented joint posterior $pr\left(\boldsymbol{\beta}, \sigma^2, v, \boldsymbol{\omega} \mid \mathbf{y}, \mathbf{X}\right)$, which the GS breaks down into consecutive draws of conditional components.

Step 1: Draw $\boldsymbol{\beta}$

Given our multivariate-normal choice of prior for $\boldsymbol{\beta}$ the conditional posterior for this vector can be derived in straightforward fashion (e.g. Lindley and Smith 1972) as:

$$pr\left(\boldsymbol{\beta} \mid \mathbf{y}, \mathbf{X}, \sigma^2, \boldsymbol{\omega}\right) = mvn\left(\boldsymbol{\mu}_1, \mathbf{V}_1\right) \qquad \text{where}$$

$$\mathbf{V}_1 = \left(\mathbf{V}_0^{-1} + \mathbf{X}'\left(\sigma^2\boldsymbol{\Omega}\right)^{-1}\mathbf{X}\right)^{-1} \text{ and } \quad \boldsymbol{\mu}_1 = \mathbf{V}_1\left(\mathbf{V}_0^{-1}\boldsymbol{\mu}_0 + \mathbf{X}'\left(\sigma^2\boldsymbol{\Omega}\right)^{-1}\mathbf{y}\right).$$

Step 2: Draw σ^2

Applying again standard results for generalized regression models, we obtain

$$pr\left(\sigma^2 \mid \mathbf{y}, \mathbf{X}, \boldsymbol{\beta}, \boldsymbol{\omega}\right) = ig\left(\eta_1, \kappa_1\right) \quad \text{with} \quad \eta_1 = \left(n + 2\eta_0\right)/2 \quad \text{and} \quad \kappa_1 = \frac{1}{2}\left(\left(\mathbf{y} - \mathbf{X}\boldsymbol{\beta}\right)' \boldsymbol{\Omega}^{-1}\left(\mathbf{y} - \mathbf{X}\boldsymbol{\beta}\right) + 2\kappa_0\right).$$

Step 3: Draw v

The relevant kernel for draws of v is its prior times the density of $\boldsymbol{\omega}$, i.e. $pr\left(v \mid \boldsymbol{\omega}\right) = \frac{1}{v_0}\exp\left(-\frac{v}{v_0}\right) \cdot \prod_{j=1}^{n} \frac{\left(\frac{v}{2}\right)^{\frac{v}{2}}}{\Gamma\left(\frac{v}{2}\right)} \omega_j^{-\left(\frac{v}{2}+1\right)} \exp\left(-\frac{v}{2\omega_j}\right)$. This is a non-standard density, and we use a random walk Metropolis-Hastings algorithm (MH, Hastings 1970 Chib and Greenberg 1995) to take draws from this kernel. Specifically, we draw a candidate value of v_c in the r^{th} round of the GS from a truncated-at-zero normal proposal density with mean v_{r-1}, i.e. the current value of v, and standard deviation s_v, and accept the draw as the new current value with probability $\alpha_v = \min\left(\frac{pr\left(v_c \mid \boldsymbol{\omega}\right)}{pr\left(v_{r-1} \mid \boldsymbol{\omega}\right)}, 1\right)$.

The standard deviation of s_v is chosen (after some trial and error in preliminary runs) to yield an acceptance probability in the 45-50% range, as suggested by Gelman et al. 2004. Ch. 11.

<u>Step 4:</u> Draw $\boldsymbol{\omega}$

For this step we note that $\frac{\varepsilon_j}{\sigma} \sim n\left(0,\omega_j\right)$. We can then use again standard results for the Gaussian regression model to obtain $pr\left(\omega_j \mid y_j, \mathbf{x_j}, \boldsymbol{\beta}, \sigma^2, v\right) = ig\left(\psi, \zeta\right)$ with $\psi = \left(v+1\right)/2$ and

$\zeta = \frac{1}{2}\left(\left(y_j - \mathbf{x}'_\mathbf{j}\boldsymbol{\beta}\right)^2 / \sigma^2 + v\right)$.

Appendix B: Specification of Refined Priors:

This appendix describes the detailed derivation of refined priors for slope coefficients in models indexed with "*b*" or "*c*" in Table 3. For ease of exposition we label the relevant components of $\boldsymbol{\beta}$ as follows:

β_{inc}	slope coefficient for log(income)
β_{lna}	slope coefficient for log(acres), where applicable
β_{lna2}	slope coefficient for $[\log(\text{acres})]^2$, where applicable
β_{a}	slope coefficient for (acres, in 000), where applicable
β_{a2}	slope coefficient for $[(\text{acres, in }000)]^2$,where applicable

Refinements for β_{inc}:

Given the log-form for the dependent variable in our MRMs β_{inc} can be interpreted as income elasticity with respect to WTP. Three of the source studies underlying our meta-dataset include a statistically significant income variable in their model. Blomquist and Whitehead 1998 regress log(WTP) against income (in \$000) and estimate a coefficient of 0.03. However, for our specification we need a prior distribution for $\frac{\partial \log(wtp)}{\partial \log(inc)}$. This requires a conversion of Blomquist and Whitehead 1998's result. Noting that in their model $wtp = \exp\left(\gamma_{inc} \cdot \left(inc/1000\right)\right)\exp\left(\text{"rest"}\right)$ we can write

$$\frac{\partial wtp}{\partial inc} = wtp \cdot \left(\gamma_{inc}/1000\right) \text{ and}$$

$$\frac{\partial wtp}{\partial inc}\frac{inc}{wtp} \approx \frac{\partial \log(wtp)}{\partial \log(inc)} = \left(\gamma_{inc}/1000\right) \cdot inc.$$

Using their sample mean of income, converted to 2006 dollars, we derive an approximated point estimate for income elasticity of 1.146. The second source study that relates WTP to income is Poor 1999. This study also estimates a model of log(wtp) on log(income), so we can directly adopt the reported coefficient of 0.12 as a second prior point estimate for β_{inc}.

The third study is Klocek 2004. The author estimates a linear bid function model relating WTP to income (in dollars). Dividing by the coefficient on "bid" we derive a scale-corrected income coefficient of 0.00022, i.e. $\frac{\partial wtp}{\partial inc} = 0.00022$. Using again sample means for income and WTP we convert this figure to an approximate point estimate for income elasticity of $\frac{\partial wtp}{\partial inc}\frac{inc}{wtp} = 1.23$.

We then treat these three estimates as draws form a normal prior distribution with mean μ_0 and variance V_0. Further, we impose the constraint of very small probability mass (say <0.01) for negative values given the vast evidence from the empirical valuation literature of non-decreasing WTP over income. We thus employ a constrained maximum likelihood routine to estimate the most likely values for μ_0 and variance V_0 given the three data points. We then use the resulting estimates of $\hat{\mu}_0 = 0.9413$ and $\hat{V}_0 = 0.4046^2$ as prior mean and variance for β_{inc} in our refined models.[10]

Refinements for $\beta_{\ln a}$

Given their focus on single wetland applications none of our source studies explicitly include a variable corresponding to wetland size in their regression model. However, most of the existing meta-analyses on wetland valuation include this regressor. For example, Woodward and Wui 2001 relate log(wtp/acre) to log(acres) with an estimated coefficient of -0.286 (model C). We can write their model as $\log\left(\frac{wtp}{acres}\right) = \gamma_{ac}\log(acres) + \text{"rest"} \xrightarrow{\text{ergo}} \log(\text{wtp}) = (1+\gamma_{ac})\log(acres) + \text{"rest"}$. It follows that $\frac{\partial \log(wtp)}{\partial \log(acres)} = (1+\gamma_{inc}) = 0.714$, which serves as our first point estimate for $\beta_{\ln a}$. Borisova-Kidder 2006 regresses log(wtp) against acres, with an estimated coefficient of 0.000000965. Following the arguments for the Blomquist and Whitehead 1998 study above we derive an expression for "acreage elasticity" as $\frac{\partial wtp}{\partial acres}\frac{acres}{wtp} = \gamma_{ac} \cdot acres$. Using the sample mean of 270,758 acres this yields a second point estimate for $\beta_{\ln a}$ *of* 0.26. Brander et al. 2006 regress log(wtp/hectares) against log(hectares) and estimate a coefficient of -0.11. Using the logic applied to Woodward and Wui 2001 this implies a point estimate of 0.89 for log(wtp) against log(hectares). This marginal effect is the same for "acres" since the conversion factor would be absorbed in the intercept. Feeding these three point estimates into our constrained ML routine described above yields estimates for the mean and variance of the refined prior distribution for $\beta_{\ln a}$ of 0.6213 and 0.2654^2, respectively.[11]

Refinements for $\beta_{\ln a2}$

Using a change-in-variable approach it is straightforward to show that if $\frac{\partial \log(y)}{\partial \log(x)} = a$,

$\frac{\partial \log(y)}{\partial (\log(x))^2} = \frac{1}{2} a \cdot (\log(x))^{-1}$. Thus, using a mean acreage of 963,466 for Woodward and Wui 2001 we obtain an approximate estimate of $\frac{\partial \log(wtp)}{\partial (\log(acres))^2} = \frac{1}{2} 0.714 \cdot (\log(963,466))^{-1} = 0.026$. The analogous values for Borisova-Kidder 2006 and Brander et al. 2006 are 0.01 (study mean = 270,758 acres) and 0.054 (study mean = 4049 acres), respectively. Unconstrained ML applied to these three point estimates produces estimated prior moments of $\hat{\mu}_0 = 0.03$ and $\hat{V}_0 = 0.0065^2$.

Refinements for β_a

We use the same three meta-studies to derive prior distributions for this coefficient. For Woodward and Wui 2001 we need to convert $\frac{\partial \log(wtp)}{\partial \log(acres)}$ into $\frac{\partial \log(wtp)}{\partial (acres/1000)}$. Using again a change-in-variable approach it can be shown that if $\frac{\partial \log(wtp)}{\partial \log(acres)} = a$, $\frac{\partial \log(wtp)}{\partial (acres/1000)} = 1000a/acres$.

Using a = 0.714 and the sample mean of 963,466 acres we obtain a first point estimate for β_a of 0.00074. For the Borisova-Kidder 2006 study we simply need to multiply the reported coefficient by 1000 to obtain our second point estimate of 0.000965. For Brander et al. 2006 we proceed as for Woodward and Wui 2001 to derive a third estimate of 0.022. Imposing again the constraint of a positive marginal effect of acreage on WTP our ML routine produces estimates of 0.0151 and 0.0065^2 for the prior mean and variance of β_a.

Coefficient for β_{a2}

Given our results for β_a mean estimates for β_{a2} would be arbitrarily close to zero. We thus retain a non-informative prior mean of 0 for this coefficient but reduce the prior variance to 0.1 in our refined specifications.

Notes

[1] We are only aware of one such study, where BT is employed to decide on critical habitat designations (Loomis 2006).

[2] Since only one study furnishes multiple observations on WTP in this application we abstract from the modeling of panel-structures and treat each observation as flowing from a different source.

[3] While this *S*-fold replication is optional it is computationally inexpensive and improves the efficiency of the predictive distribution.

[4] For a comprehensive discussion of Bayesian Model Averaging (BMA) see e.g. Raftery (1995), Hoeting et al. (1999), and Chipman et al. (2001).

[5] Naturally, one could widen this model space with additional variants of our chosen specifications. However any additional feasible model is unlikely to carry considerable posterior weight.

[6] Given our focus on model weights and posterior *predictive* distributions detailed results for posterior distributions of individual parameters for each model have been omitted from this text. They are available from the authors upon request.

[7] To guard against dramatic outliers we further truncate the *exponentiated* distribution of our logged predictions at the 99th percentile, i.e. we discard the 500 largest observations. This final adjustment is implemented in identical fashion for all models. Intuitively, this correction could be interpreted as "imposing income constraints" on the predicted WTP values.

[8] The *nse* is computed as $std/\sqrt{\left(R_p\right)}$ where *std* is the standard deviation of the predicted distribution and R_p is the number of simulated draws for the predicted series.

[9] One could argue that the SCNA and SPNA ought to be valued separately. We decided to pool the two areas for valuation purposes since this strategy best corresponds to the bulk of scenarios underlying our meta-data. In most of these studies, respondents were asked to value groups, bundles, or large areas of

non-contiguous wetlands. A separate valuation of the Spring Valley areas using our MRM would likely lead to an over-estimation of combined economic benefits. Naturally, it would be straightforward to design a primary valuation study that elicits separate benefit figures for the two areas.

[10] One might alternatively consider using reported standard errors in source studies to derive prior variances. However, standard errors (i.e. empirical variability of estimated coefficients) in a classical framework simply indicate noise resulting from sampling error, i.e. the notion of extrapolating from a finite sample to an underlying population. This concept is completely absent in Bayesian methodology, where stipulated prior distributions already correspond to the underlying population of interest.

[11] Intuitively, WTP should be non-decreasing in wetland size. However, in our case the non-negativity constraint emerged as non-binding in the ML routine.

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Table 1: Source Studies Used in Meta-Regression

study ID	Authors	Publication type	Publication year	Study year	Study Area	Underlying target population	response rate
1	Loomis et al .	journal article	1991	1989	Wetlands in the San Joaquin Valley, CA	San Joaquin Valley households	35%
2	Hanemann et al.	journal article	1991	1989	Wetlands in the San Joaquin Valley, CA	CA households outside the San Joaquin Valley	51%
3	Whitehead and Blomquist	journal article	1991	1989	Clear Creek wetland area in Western KY	Kentucky households	31%
4	D. Mullarkey	PhD dissertation	1997	1994	110 acres of undesignated wetlands in northwest Wisconsin	Wisconsin households	60%
5	Roberts and Leitch	Government Report	1997	1996	Mud Lake wetland area on SD / MN border	Households within 30 miles of study area	62%
6	Blomquist and Whitehead	journal article	1998	1990	Various wetland habitats in Western KY	Kentucky households	70%
7	J. Poor	Journal article	1999	1996	Rainwater Basin Wetlands, NB	Nebraskan households	46%
8	J. M. Tkac	Master's thesis	2002	2001	Alfred Bog, Ontario, CA	Households in the United Counties of Prescott and Russell, Ontario	57%
9	C. A. Klocek	PhD dissertation	2004	1996	Canaan Valley National Wildlife Refuge	U.S. Households	74%

Table 2: Observation-Specific Details for Meta-Regression

study ID	wetland type	original wetland area (acres)	policy scenario (acres preserved or created)	official resource designation	WTP per HH and year	HH Income	percentage of active users
1	unspecified	85,000	58,000	includes several NWRs and WMAs	284.15	66,776	46%
2	unspecified	85,000	58,000	includes several NWRs and WMAs	248.23	82,061	38%
3	bottomland hardwood forests wetlands	84,000	5,000	none	17.39	52,258	16%
4	unspecified	110	110	none	1.7 (a)	43,880 (m)	1%
5	permanently, semi-permanently, or seasonally flooded lacustrine wetlands	5,000	5,000	none	3.03	38,745 (m)	18%
6a	permanently flooded freshwater marsh	3,968	500	none	2.62	38,207	14,2%
6b	temporarily flooded bottomland hardwoods	70,080	500	none	7.27	38,207	14,2%
6c	seasonally flooded bottomland hardwoods	25,216	500	none	5.7	38,207	14,2%
6d	permanently flooded bottomland hardwood	1,408	500	none	17.37	38,207	14,2%
7	unspecified	34,000	41000 (c)	none	27.18	41,238	52%
8	domed peat bog with boreal forest	10,378	10,378	class 1 Wetland / ANSI	4.66 (a)	46,024 (m)	29%
9	high elevation moist valley	708	23,292	NWR	0.63	64,532	2%
	means:	***33,739***	***14,707***		***61.36***	***51,077***	***25%***

All monetary figures are in 2006 U.S. dollars
(a)= originally elicited as lump sum payment; annualized using a discount rate of 6%
(s) = sample mean as reported in source study, converted to 2006 dollars
(m) = census median (sample income not reported)
HH = household
NWR = National Wildlife Refuge / WMA = Wildlife Management Area / ANSI = Area of Natural and Scientific Interest

Table 3: Model Specifications

model label	error distribution	prior mean and variance for constant term	priors for slope coefficients	log (acres)	log (acres), squared	acres (in units of 1000)	acres (in units of 1000), squared
M1a	t	0, 10	diffuse	x	x		
M1b	t	0, 10	refined	x	x		
M1c	t	-50, 100	refined	x	x		
M2a	t	0, 10	diffuse	x			
M2b	t	0, 10	refined	x			
M2c	t	-50, 100	refined	x			
M3a	t	0, 10	diffuse			x	x
M3b	t	0, 10	refined			x	x
M3c	t	-50, 100	refined			x	x
M4a	t	0, 10	diffuse			x	
M4b	t	0, 10	refined			x	
M4c	t	-50, 100	refined			x	
M5a	normal	0, 10	diffuse	x	x		
M5b	normal	0, 10	refined	x	x		
M5c	normal	-50, 100	refined	x	x		
M6a	normal	0, 10	diffuse	x			
M6b	normal	0, 10	refined	x			
M6c	normal	-50, 100	refined	x			
M7a	normal	0, 10	diffuse			x	x
M7b	normal	0, 10	refined			x	x
M7c	normal	-50, 100	refined			x	x
M8a	normal	0, 10	diffuse			x	
M8b	normal	0, 10	refined			x	
M8c	normal	-50, 100	refined			x	

Table 4: Comparison of Model Fit

	t - errors		
model	log pr(y|M)	MAPE	pr(M|y)
M1a	-53.869	0.735	0.000
M2a	-50.686	0.679	0.000
M3a	-55.964	0.471	0.000
M4a	-52.002	0.814	0.000
M1b	-52.142	0.937	0.000
M2b	-50.145	0.858	0.000
M3b	-53.262	0.820	0.000
M4b	-47.565	0.742	0.000
M1c	-56.943	1.067	0.000
M2c	-55.041	1.008	0.000
M3c	-59.316	0.802	0.000
M4c	-53.453	0.756	0.000

	normal errors		
model	log pr(y|M)	MAPE	pr(M|y)
M5a	-32.679	0.728	0.002
M6a	-29.566	0.682	0.037
M7a	-35.126	0.482	0.000
M8a	-31.511	0.748	0.005
M5b	-31.045	0.866	0.009
M6b	-29.509	0.817	0.040
M7b	-32.279	0.767	0.002
M8b	-26.385	0.729	0.902
M5c	-35.882	0.965	0.000
M6c	-34.405	0.930	0.000
M7c	-38.736	0.760	0.000
M8c	-32.226	0.734	0.003

pr(y|M) = marginal likelihood
MAPE = mean absolute percent error
pr(M|y) = posterior model weight

Table 5: Comparison of BT Predictions

t - errors

	Region		NV		NV/UT	
model	mean	nse	mean	nse	mean	nse
M1a	4.150	0.046	3.372	0.039	3.477	0.040
M2a	5.643	0.066	4.094	0.049	4.039	0.047
M3a	3.357	0.018	2.753	0.015	2.810	0.016
M4a	8.177	0.085	7.586	0.083	7.623	0.083
M1b	20.505	0.293	21.284	0.332	21.726	0.339
M2b	13.704	0.172	12.349	0.160	12.721	0.166
M3b	7.911	0.070	7.802	0.069	8.091	0.075
M4b	6.635	0.060	5.655	0.056	5.604	0.052
M1c	22.554	0.353	30.670	0.502	31.503	0.543
M2c	16.522	0.230	19.170	0.269	20.414	0.296
M3c	5.917	0.055	6.694	0.066	6.815	0.066
M4c	4.972	0.041	5.332	0.047	5.261	0.045

normal errors

	Region		NV		NV/UT	
model	mean	nse	mean	nse	mean	nse
M5a	3.480	0.031	2.801	0.025	2.787	0.025
M6a	4.540	0.037	3.368	0.029	3.426	0.030
M7a	3.266	0.015	2.699	0.013	2.704	0.013
M8a	5.674	0.042	4.766	0.037	4.816	0.037
M5b	11.223	0.115	10.300	0.113	10.341	0.115
M6b	8.874	0.078	7.740	0.073	7.893	0.076
M7b	6.017	0.041	5.766	0.042	5.808	0.042
M8b	5.427	0.037	4.624	0.033	4.688	0.034
M5c	10.108	0.117	11.575	0.142	11.768	0.145
M6c	8.236	0.079	9.023	0.093	9.329	0.097
M7c	4.410	0.030	4.659	0.034	4.774	0.034
M8c	4.029	0.026	4.098	0.027	4.224	0.028

weighted average, all models

	Region		NV		NV/UT	
	mean	nse	mean	nse	mean	nse
per HH	5.577	0.040	4.749	0.035	4.817	0.036
total	62,005	442	3,567,616	26,329	6,996,222	52,042

nse = numerical standard error of the posterior mean

Toward Benefit Estimates for Conservation Programs in Agriculture – Meta Analyses for Improvements in Wetlands, Terrestrial Habitat, and Surface Water Quality

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I. Setting the Context

Conservation programs have been a significant component of the agricultural policy landscape since the 1980s. They are extensive, complex, and represent a substantial public investment, so it comes as no surprise that there is widespread interest in assessing their effects and benefits. USDA has undertaken a major Conservation Effects Assessment Program (CEAP) and, while it stops short of formal benefit assessment, USDA/ERS has begun some serious inquiries into benefits. The research reported here was undertaken in response to ERS interest in assessing the benefits of conservation programs in agriculture.[1]

Randall (2007) outlined a consistent valuation and pricing framework for conservation programs in agriculture, in which programs generate values (not directly, but via effects that modify the quantity and quality of valued services), and these values (reflecting quantity, quality, and location of services produced) are implemented at the farm level as green prices. Flury *et al.* (2005) demonstrate the welfare gains from targeting green prices to reflect regional differences in productivity of environmental services.

Here, we report three new meta analyses estimating value functions for agricultural conservation program impacts on water quality, wetlands, upland wildlife. Rather than goodness of fit, our highest priority was to obtain robust parameter estimates for key variables describing environmental services, relevant demographic variables, and regional effects, while controlling for methodological differences among studies. Our meta-analytic estimates demonstrate the ability to control for some familiar empirical effects of valuation method, and to estimate the marginal influence on value (or WTP, as the case may be) of scale and scope of service improvements, and (in some cases) demographic characteristics of the human demander populations studied. Nevertheless, we claim only that the cup is half-full – the empirical evidence is modest compared with the valuation task, and valuation studies are still designed, it seems, more for methods development than to build a generalizable empirical record.

We then focus in, for some closer analyses of our wetlands meta analysis – wetlands is an especially challenging case – that highlight the prospects and challenges for benefits transfer based on estimated meta-analytic equations.

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[1] We are grateful for research support from US Department of Agriculture, Economic Research Service under Cooperative Agreement 43-3AEL-4-80094, A Meta-analysis of Environmental Benefits from Conservation Programs in Agriculture. LeRoy Hansen served as ERS cooperator, and we appreciate his unwavering enthusiasm for the project. LeRoy Hansen, Dan Hellerstein, and others have offered helpful suggestions.

We conclude with a few comments on, first, the challenges in using meta analysis for benefits transfer and, then, the tasks that would remain – assuming we had satisfactory benefit transfer tools for environmental services – to complete a benefit assessment of conservation programs in agriculture.

The valuation context

Conservation objectives are supported by a complex web of policies and programs designed to influence farmer decisions so as to increase production of a considerable variety of environmental services farm-by-farm on a regional or national scale. In addition to the well-known challenges of valuing nonmarket services, there is the challenge of valuing not just services but programs and policies, and doing so within a framework that is sensitive to spatial considerations and consistent as we move from single to multiple amenities, and from local to national spatial scales, and back again.

Valuing conservation policies and programs. The task, familiar but nevertheless challenging to economists, of valuing non-commodity services is only the final step in the process of valuing policies and programs (Figure 1). This process involves several intervening steps – linking policies and programs to effects, linking effects to changes in services, and valuing those changes in services.[2] Programs provide incentives and perhaps constraints that, interacting with environmental conditions and farmer decisions, generate effects (e.g., additional areas of terrestrial habitat are produced). This habitat produces various environmental services (e.g., wildlife improvements that support ecosystem integrity and enhance recreational opportunities), a process mediated by farmer and user decisions. These services are valued by users and passive users. The whole process takes place within a web of public laws, policies, and regulations.

Specifying the transformations from programs to effects[3] and from effects to services would seem primarily a task for natural science, while valuing services seems primarily the responsibility of economists. Yet effects, services, and value all emerge from the interaction of natural and economic systems, and can best be understood through interdisciplinary collaboration among natural scientists and economists.

Spatial considerations, and complex policy. At every step in the process from policies and programs to changes in the level of environmental services, spatial considerations matter: spatial scale and scope, border effects, and pattern effects produced by various discontinuities and nonlinearities. For valuing changes in environmental services, scarcity and substitution/complementarity relationships are, among other things, spatial in nature and systematically affect nonmarket values (Schläpfer and Hanley 2003).

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[2] While the arrangement illustrated in the Figure (effects generate services, which are valued) is typical, valuation efforts may sometimes be addressed more readily to effects than to services. People value habitat for its various services that they use actively or passively, yet it may be convenient to enquire about their WTP for habitat directly, rather than to specify its services and value each of them. In other cases, an extra step prior to valuation is involved – habitat enhances wildlife which supports increased wildlife viewing activity, and these additional activity-days are valued.

[3] There exists a substantial scientific literature on effects. For example the US Department of Agriculture's Conservation Effects Assessment Program has compiled a 454-item bibliography (Gagnon et al. 2004).

From a valuation perspective, conservation programs constitute a complex policy, in the sense of Hoehn and Randall (1989), who show the default valuation procedure (independent piecewise valuation) is generally invalid[4], whereas a valid valuation scheme for complex policies is theoretically and empirically much more demanding. To avoid the independent piecewise valuation problem, the outputs of multifunctional agriculture should be valued as a package on a national scale[5].

Meta-analysis: generalizing from the valuation studies literature

There is a large literature reporting environmental valuation studies worldwide. A complete compendium is elusive, but the magnitude of some incomplete lists is impressive. Carson (2003) reports more than 5,000 studies using contingent valuation methods. Navrud and Vagnes (2000) identified 650 European studies, and one would expect the number to be considerably greater today. Richard Bishop (personal communication) maintains a bibliography of wetlands valuation studies that has grown to about 400 entries. Villa *et al.* (2002) list 146 ecosystem valuation studies worldwide. Environmental valuation data bases listed by Genty (2005) contain a total of more than 2,350 studies, although some duplication across the different data bases seems inevitable.

New studies will continue to be desirable, if relatively expensive, and should be done to further develop and test methods, and to augment, update, and improve the body of empirical evidence. Nevertheless, it is inevitable that attention be addressed also to generalizing from the substantial body of existing valuation research.

Meta-analysis has become the standard method of searching for general patterns in a body of existing specific research results (Hedges and Olkin 1985, Hedges 1992, and Lipsey and Wilson 2001). Borisova-Kidder (2006) has identified 28 completed meta-analyses of environmental services. Representative studies include Smith and Huang (1995) on air pollution, Dalhuisen (2003) on residential water demand, and Rosenberger and Loomis (2001) on outdoor recreation. A general model of the following type is estimated with regression techniques:

$$WTP_{i,j,k,l} = f(\Delta\ Service_{j,k}\ ,\ Subst/com_{\ j,k}\ ,\ Demographic_{\ i,k}\ ,\ Research\ procedure_{\ j,k,l}\ ),$$

where the four categories of independent variables are expressed as vectors, and *i* : person or household; *j* : service type; *k* : location; and *l* : valuation project.

WTP per capita (or per household) is hypothesized to be influenced by the change in level(s) of environmental service(s), the availability of substitute services, relevant demographic variables, and the research procedures used in value estimation. For

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[4] Hoehn and Randall (1989) show that, as the number of components in a complex policy expands, substitution and complementarity relationships proliferate and the budget constraint becomes more pressing. In general, WTP for a *j*-component policy $\neq \Sigma\ WTP_j$ valued independently. As j becomes large, in the limit, WTP for a *j*-component policy becomes strictly less than $\Sigma\ WTP_j$.

[5] Hoehn (1991) and Hoehn and Loomis (1993) have made progress toward a practicable framework for consistent valuation of complex policy.

meta-analysis, each study constitutes a single observation (if it reports a single valuation) or a single panel of observations if it reports valuations of, say, several options that vary in scale and scope of environmental improvements. To enjoy a reasonable prospect of success, a meta-analysis project requires a sufficiently large set of independent studies, each with methods and results reported in sufficient detail, and all sharing at least a degree of methodological consistency.

Economists, responding to the extensive data requirements of meta-analysis, have assembled environmental valuation data bases for that purpose. Table I lists three internet-searchable data bases that have potential application to environmental services of conservation programs (EVRI, Envalue, RED) and two additional data bases (RUVD and ValueBase) that, while not internet-searchable, are readily transformable for meta-analysis. The total number of studies included exceeds 1,800 (although, again, some duplicate entries are likely). However, the existing empirical literature is not quite so rich as may at first appear. When we start eliminating studies for various good reasons – some address amenities unrelated to agriculture, some do not report sufficient information about research procedures to enable independent assessment of their validity, some provide no evidence of peer review, etc. – the numbers diminish markedly. Nevertheless, a substantial body of usable studies remains.

But, even here, there is less that meets the eye. The methodological norms of the economics discipline place high priority on originality, novelty, and innovation in theory and methods, and correspondingly low priority on accumulating generalizable empirical evidence from primary data. Meta-analysis is hindered in two ways – the basic details of data gathering, handling, and analysis often are reported inadequately; and, while the use of standard protocols would advance the quest for empirical knowledge, the publication norms in economics (originality is prized, and rejection rates are high) serve to discourage their use. As a result, meta-analyses of economic valuation studies typically must deal with data sets that are small given the task at hand, and relatively noisy.

Benefits transfer

Benefits transfer (BT) seeks to economize on valuation research costs by applying the findings of particular local valuation studies to a broader set of sites (Bergstrom and de Civita, 1999, Smith *et al.*, 2001, Van den Bergh and Button, 1999). BT may take various forms. In its simplest configuration, benefits estimated at one site are applied (with only *ad hoc* modifications) to illuminate policy options at another site. Unfortunately, empirical tests of simple BT models have not yet vindicated the decision-makers' enthusiasm for the savings in research costs that BT promises (Navrud and Ready 2006). A more sophisticated approach is based on meta-analysis. Benefit estimates are obtained for a policy site by plugging policy-site-specific values for right-hand-side variables into an estimated meta-analytic equation. Assuming the meta-analytic equation is reasonably robust, this approach is preferred because it replaces the *ad hoc* adjustments of the simple approach with estimated effects generalized from the inventory of empirical studies that pass some tests of quality and relevance. However, Genty (2005) warns that estimated meta-analytic equations of environmental service values tend to be unbiased but imprecise, which suggests that application to particular policy sites may be hampered by wide confidence limits.

II. Meta Analyses for Improvements in Wetlands, Terrestrial Habitat, and Surface Water Quality

We approached the meta analysis task with a clear objective, benefits transfer. For BT, the appropriate focus is on obtaining robust estimates of how environmental services (type, quantity and quality – baseline and change) and regional factors[6] impact valuations, while controlling for differences in methodological differences among valuation studies. The available data is incomplete and inconsistent – original studies use a variety of methods, vary in quality, and offer only spot coverage of the range of service types, quality levels, and regions relevant for benefit cost analysis. So, rather than placing highest priority on goodness of fit (which at worst has the analyst chasing data points all over the map by proliferating dummy and interaction variables, even at the risk of estimating study-specific variables), we sought robust estimates of parameters useful in benefit transfer.

We considered three categories of environmental services – wetlands, terrestrial habitat, and surface water quality.

Wetlands. We started with the set of studies assembled by Woodward and Wui (2000), conducted our own search for additional studies, and applied our own selection criteria, eventually settling on a set of 72 valuations from 34 US studies, which are listed in Borisova-Kidder (2006). Variables used in our meta analysis are listed in Table 2.

Terrestrial habitat. We assembled a set of 23 valuations from 12 US studies, which are listed in Borisova-Kidder (2006). Variables used in our meta analysis are listed in Table 3.

Surface water quality. We assembled a data set from scratch, and then were able to augment it with the data set of Johnston *et al.* (2005). After applying our own selection criteria, we eventually settled on a are listed in Borisova-Kidder (2006). Variables used in our meta analysis are listed in Table set of 98 valuations (total value, i.e., use and nonuse) from 40 US studies, which 4.

While the studies chosen include some that provide single value observations and some providing panels of observations, we tested for, and rejected, fixed and random events in each case. Accordingly, we settled on OLS estimation. Our estimation results are presented in Tables 5 (wetlands), 6 (habitat), and 7 (water quality). In each case, the log-linear specification was chosen, as is often the case in meta analysis of environmental service values. Log-linear specifications typically produce estimates that are robust around the mid-range of the data, but are less plausible near and beyond the endpoints of the data series (implications for BT are highlighted in the next section).

Summary of results. Our meta-analytic estimates demonstrate the ability to control for some familiar empirical effects of methodological differences among studies, and to estimate the marginal influence on value (or WTP, as the case may be) of scale and

[6] Regions are defined in terms of the ERS Farm Resources Regions (Figure 2).

scope of service improvements, and (in some cases) demographic characteristics of the human demander populations studied.

For wetlands and terrestrial habitat, where several different valuation methods were used, the dependent variable was value/acre. For wetlands, the scale effect is small and negative but insignificant, the amount of wetlands in the vicinity (potential substitutes) is negative but not quite significant, and income is positive and significant (Table 5, Model I). Some methodological variables (as well as some wetlands service types and regions) are significant. Applying our estimated meta-equation with variables set near the mid-range of the data, we estimate that the public's WTP to enroll wetlands into the Wetlands Reserve Program is about $250/acre/year, with adjustments for wetlands type, services, and geographic regions. For terrestrial habitat, the scale effect is large and positive but insignificant, while the provision of viewing and open space services has positive and significant impact on value/acre (Table 6). For both wetlands and terrestrial habitat, more recent studies generated significantly higher values, suggesting that public appreciation of wetlands and terrestrial habitat is increasing. Estimated WTP to maintain 100 acres of mature habitat in the Conservation Reserve Program is about $130/acre/year, with adjustments for habitat services provided. We note, however, that the habitat meta-equation was based on relatively few studies, and that results seem overly sensitive to the year and number of acres.

Given that all studies in the water quality data set used contingent valuation, the dependent variable for water quality was WTP/household. The extent of the quality improvement and the size of the water body affected have small, positive and significant impact on WTP, income is not significant, some methodological variables are significant, and WTP for improvements to saltwater bodies is greater than for freshwater bodies (Table 7). More recent studies generated modestly but significantly lower WTP/household, perhaps reflecting methodological adjustments to correct perceived upward bias in contingent valuation. Estimated WTP to improve 100 miles of freshwater-body (linear miles for streams, and perimeter miles for lakes) from "suitable for rough fishing" to "swimmable" is about $50/household annually. The additional WTP/household is modest as waterbody size increases, but the number of households who would pay is typically greater for larger waterbodies (in other words, our results are consistent with the view of water quality as a local or regional public good).

These meta-analytical results may be compared with results from other recent studies. For wetlands, we were able to estimate the influence on value of four wetlands types and four regions of the US; Woodward and Wui (2000) did not examine the influence of region, and considered only two wetland types. Brander *et al.* (2003), using a global data set with three times as many observations as our US data set, were able to estimate the influence of a broader range of wetland types – additional observations are always helpful, but we were limited by our objective to estimate wetland values for the US. The water quality case illustrates the point made earlier, about the trade-off between goodness of fit and usefulness for benefits transfer: Johnston et al. (2005), using a data set that overlaps ours to a considerable extent, obtained better goodness of fit than we did, but we would argue that our estimates are more robust for benefits transfer.

Despite the fairly high noise factor in data sets of valuation studies, some clear signals have emerged from this work. Successes include the ability to control for some familiar empirical effects of methodological differences among studies, and to estimate the marginal influence on WTP of scale and scope of service improvements, and (in some cases) demographic characteristics of the human demander populations studied. For wetlands, the size of the wetland has positive and significant impact on WTP while the scale effect is negative and insignificant, the amount of wetlands in the vicinity (potential substitutes) is negative and not quite significant, and income is positive and significant. A considerable variety of market-based, behavioral, and stated preference methods are represented in the data, and some methodological variables (as well as some wetlands service types and regions) were significant. For terrestrial habitat, the area of land affected had positive and significant impact on WTP, as did the provision of viewing and open space services. For water quality, the extent of the quality improvement and the size of the water body affected had positive and significant impact on WTP, income was positive but not significant, some methodological variables were significant, and WTP for improvements to saltwater bodies was greater than for freshwater bodies.

III. A Closer Look at the Wetlands Meta Analysis

Wetlands valuation provides special challenges for meta analysis in support of benefits transfer. First, the data set is very thin, compared to the task at hand. We specified 4 types of wetlands, ERS is interested in 9 Farm Resource Regions, and our 72 observations from 34 unique studies used 8 distinct valuation methods. Since wetland type, region, and method are composed of mutually exclusive categories, these three dimensions generate a value matrix with 288 cells. Obviously, at least 216 (i.e., 288 - 72) of those cells must be empty. But it gets worse: we also identified 8 distinct environmental services that wetlands might provide. Most of these are not mutually exclusive, which implies that there are literally thousands of unique combinations of type, services, regions, and valuation methods, yet we can call upon a mere 72 observations from 34 studies.

Second, the eight distinct valuation methods represented vary widely in their theoretical properties. Energy analysis pays little attention to economic foundations; replacement cost is driven by the supply-side with no attention to evidence as to whether replacement would be demanded at cost; the net factor income, production function, and marginal productivity methods are basically “back of the envelope” budgeting methods that capture at best only a subset of wetlands service values; hedonic price analysis captures only a subset of wetlands service values, and is applied selectively to wetlands surrounded by developed property; and the travel cost method captures only a subset of wetlands service values, and is applied selectively to wetlands that attract substantial visitation. Only the contingent valuation method is demand-based, applied to wetlands of all kinds, and captures the gamut of active and passive use values – but CV has its own well-documented challenges. It follows that persistent differences in performance among these eight valuation methods should be expected.

Finally, examination of the data identifies an empirical challenge: the value estimates, standardized to $/acre/year, cover a vast range from less than $1 to more the $1million.

This range introduces a lot of variation into the data set; at a more intuitive level, the extreme observations extend credibility to the breaking point.

Given these challenges, we now explore some issues in choosing a model and setting appropriate scenarios for benefits transfer.

Choosing a model for benefits transfer

How far to pursue goodness of fit? The model with the best fit does not necessarily generate the most plausible value estimates for benefits transfer. In pursuit of generalizability, we compressed the 9 ERS Farm Resource Regions into 4: R1 includes Northern Crescent and Northern Great Plains; R2 includes Fruitful Rim and Southern Seaboard; R3 includes Heartland and Mississippi Portal; and R4 includes Eastern Uplands, Prairie Gateway, and Basin and Range. Also, we minimized the use of dummy variables and, especially, interaction terms, to minimize the risk of estimating study-specific coefficients. Applying these strategies, we sacrifice a little goodness of fit in order to get more plausible value estimates. To put it another way, a little smoothing does not hurt when one is dealing with data sets that exhibit implausibly large variation.

What to do about the two energy analysis observations? The two energy analysis studies generated two of the three highest value observations in the data set, and the mean value using EA is two orders of magnitude greater than the mean value using any other method. Does it therefore make sense to drop the two EA observations? Comparing model I with model II (Table 5) provides little justification for dropping the EA observations. Given that model II has lower R^2 and fewer significant variables, it is not clear that it should be the preferred model. Given the implausibility of the EA values, how could this be? We hypothesize that, given the thinness of the data set compared to the BT task, the EA observations contain information – for example, while they clearly overstate the values of the wetlands where they were applied, it seems likely that these were in fact relatively high-valued wetlands (just not so valuable as EA would suggest).

Suppose we are interested in value calculations from a particular valuation method**.** Despite the challenges involved, contingent valuation has some clear advantages over other valuation methods used in our data set: it is applicable to wetlands of all kinds and it captures all kinds of active and passive use values. So, an analyst may want to conduct BT as though all values had been estimated with CV. Should the analyst (1) re-estimate the meta analysis using just the 28 CV observations, or is it a better strategy to (2) use the 72 observations data set and simply set CV = 1 and all other methods at zero when calculating BT values?

Given the thinness of our data set compared to the BT task, it seems clear that the second strategy is to be preferred. The array of BT values calculated from the full model **(**Figure 3b) is much more plausible than that calculated from a model estimated with the partitioned 28 observation data set (Figure 3a). The reason seems to be that the full data set provides much better coverage of wetland types and regions and that the other valuation methods, even if the analyst considers them all inferior to CV, generate useful (if not always precise and unbiased) information about values.

Setting scenarios for value calculations

The log-linear model generates plausible estimates of the dependent variable (lnvalue) for RHS variable values near the mid-range but not, notoriously, for RHS variable values near or beyond the endpoints of their range. This effect can be seen in Table 8: values in the first block (calculated for 2007, a year that today's policy makers might find appropriate) are orders of magnitude greater than values calculated for a mid-range year (the second block). This effect applies also to the environmental service dummies. Calculated values are smaller, and more plausible (we think), when the dummies are set at their means-for-type (the third block of calculated values). There is some intuitive support for setting the service dummies at mean-for-type. Services are coded 1 when the original studies make a point of highlighting them as strengths of the wetland to be valued, and zero otherwise; and a typical study highlights just a few wetlands services, rather than the broader array one would expect to be present at some level. Therefore, it seems that mean-for-type provides a better picture of the services one could expect from a typical wetland of that type.

So, what are our best-judgment BT values calculated from our meta-analytic models?

The best-judgment BT values we can calculate from our meta-analytic models are presented in Table 9. We obtain these values from model I (Table 5), with Year set at the mid-range, Publish set at its mean, and service dummies set at mean-for-type. We offer two sets of values, one assuming methods PFMPNFI, CV, and RC in proportion to their relative frequency in the data set, and the other assuming CV only. Note that CV only gets smaller BT values.

We claim only that these BT values are the best we can calculate from our estimated meta analysis of wetlands values. Inspection and intuition suggest that not all of these values are plausible. In particular, we think the calculated values for prairie potholes are too low, and we suspect saltwater marshes are undervalued, too. Also, we wonder whether the calculated values assign too much of a premium to wetlands in Region 1 (Northern Crescent and Northern Great Plains) or conversely discount the other regions too much

Given the problems with our wetlands data set identified at the beginning of this section, we would use BT values calculated from our meta analysis as only a starting point for benefits transfer – a role remains for judgment factors beyond the estimated meta analyses.

IV. Concluding comments

Meta analysis for benefits transfer

We have demonstrated that the glass is half-full. We were able to identify systematic components of wetland value per acre, terrestrial habitat value per acre, and WTP for surface water quality improvement; and we generated a body of value estimates that provide a sound starting point for benefits transfer on a national scale.

However, the glass remains half-empty, too. After 35 years of focus on methods development, valuation research still (it seems) places a relatively low priority on building a body of generalizable evidence. Too many studies fail to meet minimal standards for inclusion in meta analysis and, among those that do, there is too little consistency in methodological details and the specification of environmental descriptors – these are serious impediments to empirical generalization. The problem is exacerbated when the needs of policy analysis proliferate RHS variables in meta analyses conducted in support of benefits transfer. For all these reasons, meta analysis of environmental values for BT is performed mostly under data-poor conditions.

If we are to move beyond the methods development stage in non-market valuation, we will need to accept the need for some standard guidance for the conduct of valuation studies, as well as for policy applications of meta-analysis and benefit transfer.

<u>Conclusions from our closer analysis of the wetlands case</u>. The log-linear model is widely used in meta analysis, and it often provides a good fit to the data. However, value calculations for benefits transfer often require RHS variable values near the mid-range (which can be disconcerting to policy practitioners – "why should I set Year at 1985 when I am assessing a project which if implemented would begin to deliver services in 2010?").

The pursuit of goodness of fit can be counterproductive for BT. When trying to get the most out of data sets that exhibit rather too much variation, it seems foolish to chase data points all over the map – under those circumstances, we could use a little smoothing! Smoothing is promoted by combining categories (e.g., methods, regions) to increase the number of observations in each category, and basing estimates for particular methods on the full data set and the estimated coefficient for the method, rather than partitioning the data set by methods and making a separate estimate for each method.

Estimating the benefits of conservation programs in agriculture

Suppose we had satisfactory value estimates for a broad array of environmental services, adjustable for type, quantity, quality, and region. That would be an excellent start, but no more than that, for estimating the benefits of conservation programs in agriculture. There would still remain the challenges of relating service values to environmental effects and ultimately to policy benefits, and doing so in ways that are sensitive to spatial considerations and the consistency requirements for evaluating complex policy and scaling it up to the national level and down again. Ultimately, coherent and effective agro-environmental policy must hit the ground as green prices (reflecting quantity, quality, and location of services produced) at the farm level.

Figures and Tables

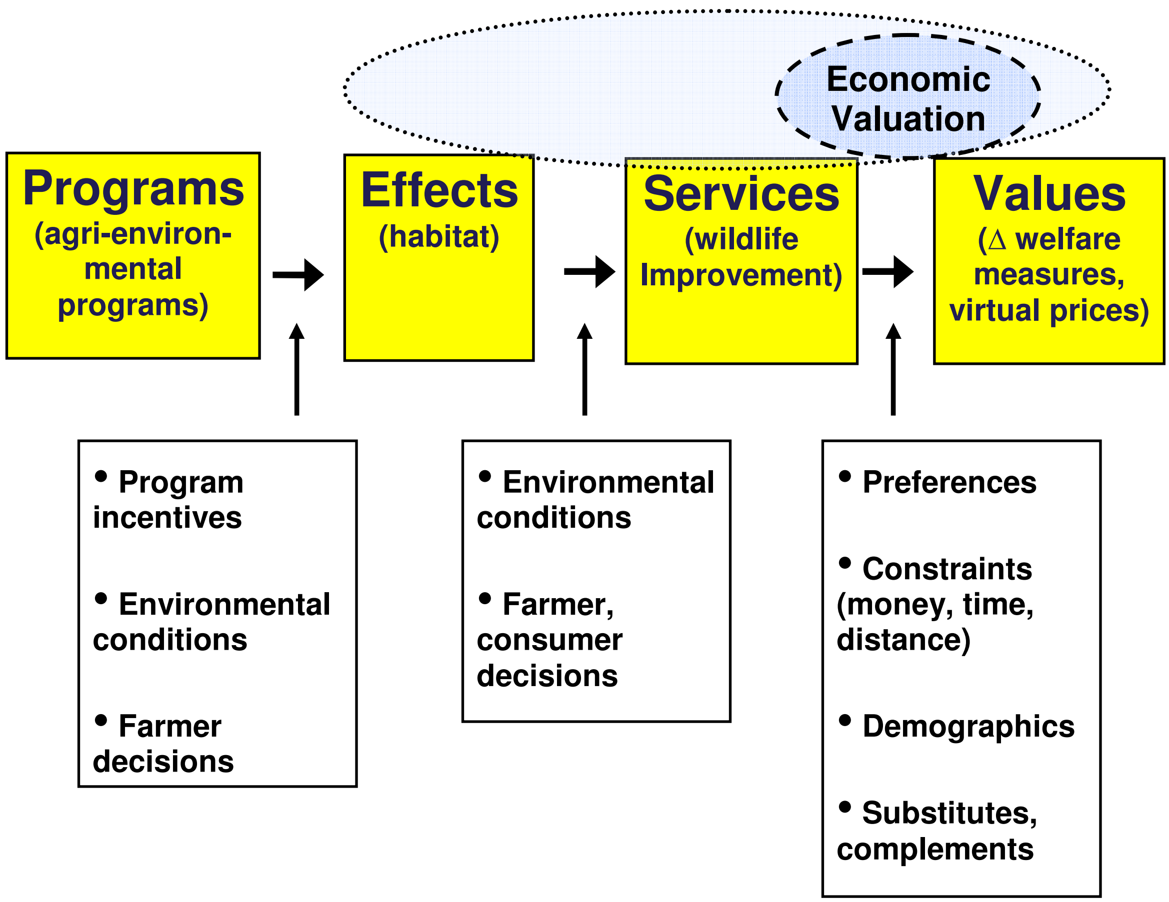


Figure 1. A framework for valuing policies and programs (Randall 2007)

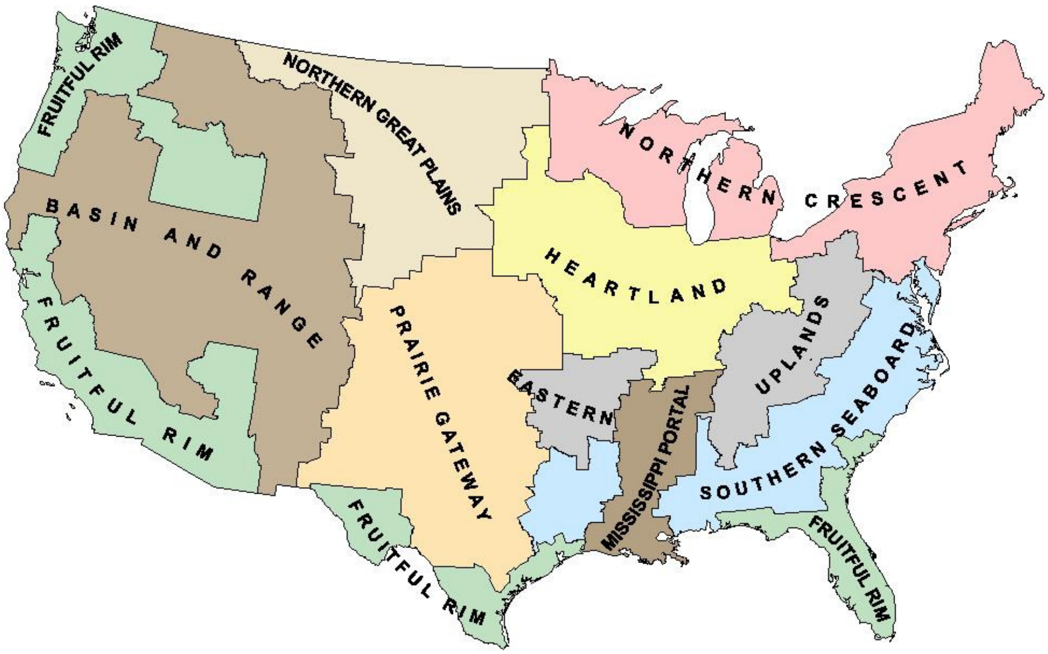
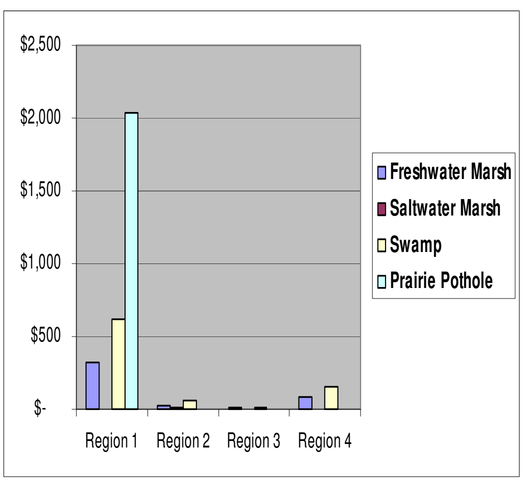
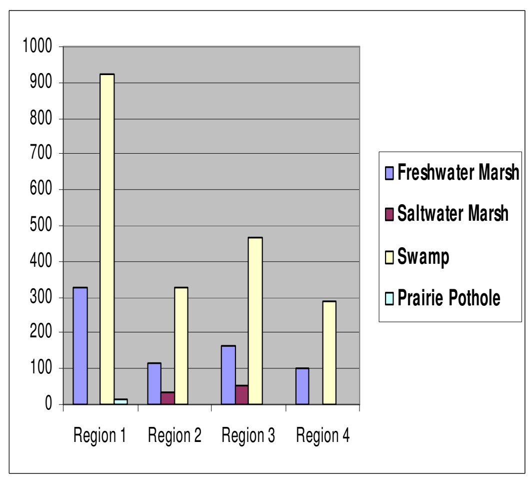


Figure 2. ERS Farm Resources Regions



(a) Partitioned data set – CV only
28 obs; R^2adj: .53; F: 2.49



(b) Full data set; method set CV = 1
72 obs; R^2adj: .53; F: 4.18

Figure 3. Calculated values assuming contingent valuation method: CV-only data set (a) and full data set (b)

Table 1. Five data sets assembled to support meta analysis of environmental services

Database	Reference, Country(ies)	Subject	Countries	Studies (valuations)
EVRI	De Civita et al. (1998), Canada	Environment, health	World	1028 (n.r.*)
Envalue	James et al. (2004), Australia	Environment, health	World	387 (n.r.*)
RUVD	Kaval and Loomis (2003), USA	Recreation	World	209 (1239)
RED	European Comm. (2003), EU	Externalities	EU	96 (n.r.*)
ValueBase	Sundberg and Soderqvist (2004), Sweden	Environment, health	Sweden	96 (410)

* n.r.: not reported.
Source: Genty (2005).

Table 2. Description of wetlands variables

Variable	Description	Frequency	Mean
Dependent variable			
LNVALUE	Logarithm of value per acre of wetland, U.S. year 2003 dollars	72	5.608
Socio-economic variables			
INCOME	Annual household income, U.S.	72	43.950
YEAR	Year in which study was conducted, 1969=1	72	16.319
Wetland size			
ACRES	No. of wetland acres (,000) valued	72	356.640
SHARE	Share of wetland acres in the area by FIPS codes as reported by the NRI 1997 data	72	0.133
Wetland types			
FRESHWATER MARSH	1 if a freshwater marsh, 0 if not	39	0.542
SALTWATER MARSH	1 if a saltwater marsh, 0 if not	19	0.264
SWAMP	1 if a swamp, 0 if not	7	0.097
PRAIRIE POTHOLE	1 if a prairie pothole, 0 if not	7	0.097
Wetland functions			
FLOOD	1 if flood reduction, 0 if not	18	0.250
WATER QUALITY	1 if water quality improvement, 0 if not	20	0.278
WATER SUPPLY	1 if water supply augmented, 0 if not	14	0.205
RECFISH	1 if recreational fisheries improved, 0 if not	23	0.319
COMFISH	1 if commercial fisheries improved, 0 if not	20	0.278
BIRDHUNT	1 if bird/wildlife hunting, 0 if not	23	0.319
BIRDWATCH	1 if bird/wildlife hunting observation, 0 if not	17	0.236
AMENITY	1 if amenities augmented, 0 if not	14	0.194
HABITAT	1 if habitat is augmented, 0 if not	23	0.319
Methodological variables			
CVM	1 if study used Contingent Valuation Method, 0 if not	28	0.389
HP	1 if study used Hedonic Pricing Method, 0 if not	3	0.042
TCM	1 if study used Travel Cost Method, 0 if not	4	0.056
RC	1 if study used Replacement Cost Method, 0 if not	16	0.222
PFMPNFI	1 if study used Production Function or Market Prices or Net Factor Income Method, 0 if not	19	0.264
EA	1 if study used Energy Analysis Method, 0 if not	2	0.028

PUBLISH	1 is study is a journal article, 0 if not	50	0.694
Regions			
R1	1 if study conducted in Northern crescent or Northern great plains, 0 if not	28	0.389
R2	1 if study conducted in Fruitful rim or Southern seaboard, 0 if not	22	0.306
R3	1 if study conducted in Heartland or Mississippi portal, 0 if not	17	0.236
R4	1 if study conducted in Prairie gateway=1 or Eastern uplands, 0 if not	5	0.069

Table 3. Description of habitat variables

Variable	Description	Frequency	Number of studies	Mean
Dependent variable				
LNBENPACRE	Logarithm of benefit per acre, U.S. year 2003 dollars	23	11	4.87
Study characteristics				
YEAR	Year study was conducted, 1982=1	23	11	9.26
LNACRE	Log number of acres valued	23	11	10.27
CVM	1 if contingent valuation method, 0 if not	21	10	0.91
PUBLISH	1 if study in refereed journal, 0 if not	19	9	0.83
Services				
VIEWING	1 if noted in the study, 0 if not	14	6	0.61
OS (open space)	1 if noted in the study, 0 if not	6	3	0.26
OSHABSING (OS + habitat for single sp.)	1 if noted in the study, 0 if not	2	2	0.09
OSHABMULT (OS + habitat for mult spp.)	1 if noted in the study, 0 if not	9	5	0.39

Table 4. Description of water quality variables

Variable	Description	Frequency	No. of studies	Mean
Dependent variable				
LNWTP	Log WTP for surface water quality improvements per household per year, 2003 dollars	98	40	4.63
Surveyed population				
INCOME	Annual household income, 2003 dollars	98	40	48162.
NONUSERS	1 if nonusers sample was used in the survey, 0 if not	19	11	0.19
Methodological variables				
YEAR	Year study was conducted, 1982=1,	98	40	15.28
PUBLISH	1 if in refereed journal, 0 if not	56	21	0.57
VOLUNTCONTR	1 if voluntary contribution, 0 if not	11	6	0.11
LUMPSUM	1 if a single lump sum payment), 0 if not	74	26	0.76
MEDIANWTP	1 if value was reported as median WTP , 0 if not	5	4	0.05
NONPARAMETRIC	1 if nonparametric estimation of WTP, 0 if not	41	13	0.42
PROTESTBIDS	1 if protest bids were excluded, 0 if not	46	18	0.47
OUTLIERBIDS	1 if outlier bids were excluded, 0 if not	26	13	0.27
HIRESP	1 if response rate higher than 74%, 0 if not	32	10	0.33
Method of eliciting WTP values				
DISCRETECHOICE	1 if discrete choice method used, 0 if not	11	3	0.11
Method of survey administration				
MAIL	1 if mail method, 0 if not	37	20	0.38
PHONE	1 if phone method, 0 if not	21	6	0.21
INTERVIEW	1 personal interview method, 0 if not	28	9	0.29
MULTMETH	1 if multiple methods, 0 if not	12	5	0.12
Waterbody type, size and scale of improvement				
FRESH	1 if freshwater, 0 if not	82	30	0.84

OTHERTYPE	1 if an estuary or saltpond, 0 if not	16	10	0.16
WATERSIZE	Water body size based on hydrological and USDA/ ERS data (river, lake and coastal lines in miles)	98	40	41547.
WQCHANGE	Change in water quality (RFF water quality ladder).	98	40	2.61
WQLADDER	1 if Water Quality Ladder used in elicitation, 0 if not	40	12	0.41
Region				
MULTREG	1 if a study was conducted in multiple regions, 0 if not	32	10	0.327
R1	1 if Northern crescent or Northern great plains, 0 if not	18	11	0.184
R2	1 if Fruitful rim or Southern seaboard, 0 if not	23	10	0.235
R3	1 if Heartland or Mississippi Portal, 0 if not	20	6	0.204
R4	1 if Basin & Range, Prairie Gateway, or Eastern Uplands, 0 if not	5	3	0.050

Table 5. Wetland meta-analytical estimates, (I) 72 observations; and (II) 70 observations: energy analysis excluded

Variable	Model I (72 obs)	Model II (70 obs, EA excluded)
INTERCEPT	-3.183 (2.67)	-3.412 (2.63)
INCOME	0.123** (.056)	0.140** (.056)
YEAR	0.164*** (.057)	0.151** (.057)
ACRES	-2.507E-4 (3.58E-4)	4.672E-6 (3.00E-4)
SHARE	-5.031 (3.69)	-5.703(3.66)
FRESHWATERMARSH	-0.249 (1.10)	-0.216 (1.08)
SALTWATERMARSH	-2.173* (1.21)	-1.968* (1.20)
PRAIRIEPOTHOLE	-2.772*(1.50)	-2.873* (1.48)
WATERSUPPLY	0.442 (.761)	0.622 (.965)
QUALITY	1.858** (.640)	2.163** (.773)
FLOOD	-0.024 (.629)	0.243 (.653)
RECFISH	0.802* (.950)	0.990 (.631)
COMFISH	1.853* (.950)	1.40 ((.977)
BIRDHUNT	-0.747 (.685)	-0.756 (.674)
BIRDWATCH	2.133** (.857)	2.185** (.844)
AMENITY	-2.021* (1.84)	-1.944* (1.07)
HABITAT	-0.330 (.781)	-0.412 (.770)
PUBLISH	2.064* (1.23)	1.774 (1.22)
EA	5.132*** (.976)	
PFMPNFI	-0.701 (.918)	-0.882 (.968)
CVM	-1.739* (1.67)	-1.920** (.911)
HP	0.966 (1.23)	1.124 (1.65)
TCM	-0.565 (1.72)	-0.656 (1.22)
R1	1.163 (1.64)	0.787 (1.63)
R2	0.128 (1.43)	0.224 (1.41)
R3	0.475 (1.42)	0.324 (1.40)
K (Number of independent variables)	25	24
N (Number of Observations)	72	70
R^2 (Adj-R^2)	0.694 (0.528)	0.651 (0.465)
F	4.18***	3.49***
Durbin-Watson	1.879	1.819

*: Significantly different from zero at the 10% level
**: Significantly different from zero at the 5% level
***: Significantly different from zero at the 1% level

Table 6. Habitat estimates

Variables	Coefficients (standard errors in parentheses)
(Constant)	-10.366 (6.238)
YEAR	.465** (0.185)
LNACRE	.344 (0.369)
VIEWING	6.669*** (2.059)
OS	5.331** (2.073)
OSHABMULT	2.014 (1.555)
PUBLISH	-.272 (2.09)
CVM	1.514 (2.262)

N=23
R^2=0.583 (adjusted=0.388)
F=2.99**
Durbin-Watson=2.153

Table 7. Water Quality estimates

Variable	Model without regional variables	Model with regional variables
(CONSTANT)	5.231*** (.570)	5.856 *** (.719)
WATERSIZE	2.82E-006*** (.000)	3.16E-006*** (.000)
WQCHANGE	.128** (.054)	.118** (.056)
WQLADDER	-.224 (.177)	-.637** (.263)
INCOME	-2.60E-006 (.000)	-2.84E-006 (.000)
NONUSERS	-.261 (.174)	-.259 (.177)
PROTESTBIDS	.525*** (.175)	.515** (.233)
OUTLIERBIDS	-.424** (.194)	-.285 (.223)
NONPAR	-.326* (.190)	-.238 (.211)
LUMPSUM	.105 (.174)	.149 (.195)
VOLUNTCONTR	-.598** (.256)	-.429 (.271)
MEDIANWTP	.065 (.324)	-.232 (.360)
HIRESP	-.345* (.175)	-.361** (.186)
YEAR	-.023* (.013)	-.042** (.016)
PUBLISH	.080 (.159)	.060 (.186)
DISCRETECHOICE	.288 (.265)	.487 (.326)
MAIL	-.059 (.225)	.129 (.266)
INTERVIEW	.535** (.241)	.710*** (.265)
MULTMETH	-.221 (.281)	-.003 (.323)
FRESH	-.522*** (.177)	-.650*** (.214)
MULTREG		-.537 (.450)
R1		-.342 (.386)
R2		-.507 (.420)
R3		-.088 (.420)
N	98	98
R^2 (R^2-adjusted)	.557 (.449)	.585 (.455)
F	5.152***	4.530***
Dublin-Watson	1.554	1.686

Table 8. Benefits transfer from wetlands meta analysis: scenario choice has a big impact on BT values ($/acre/year)

	Region1	Region 2	Region 3	Region 4
Final model, 72 obs. Service dummies and Publish set at 1, year at 2007				
Freshwater Marsh; all exc ComFish	344,157	28,899	148,855	28,899
Saltwater Marsh; all exc WS, WQ, Amen, Hab		52,751	74,163	
Final model, 72 obs. Service dummies and Publish set at 1, year at mean (1985)				
Freshwater Marsh; all services exc ComFish	5,023	422	2,173	422
Saltwater Marsh; all services exc, WS, WQ, Amen, Hab		1,441	2,038	
Final model, 72 obs. Publish = 1, service dummies set at mean for type, year 1985				
Freshwater Marsh	1,342	206	581	113
Saltwater Marsh		58	202	

(Calculations assume methods PFMPNFI, CV, RC, in proportion to their relative frequency in data set)

Table 9. Wetlands value estimates ($/acre/year) calculated from final OLS model with best-judgment scenarios

Wetland Type	Region 1	Region 2	Region 3	Region 4
Methods - PFMPNFI,CV,RC				
Freshwater Marsh	704	250	353	220
Saltwater Marsh		77	106	
Swamp	1,985	705	997	620
Prairie Pothole	16			
Methods - CV only				
Freshwater Marsh	328	117	165	103
Saltwater Marsh		36	51	
Swamp	925	329	465	289
Prairie Pothole	14			

(Calculations based on final model, Year = mean, publish = mean, service dummies set at mean for type)

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PART 2: Papers Supporting Objective: "Estimate the Economic Value of Changing Recreational Access for Motorized and Non-Motorized Recreation"

IDENTIFYING DEMAND PARAMETERS IN THE PRESENCE OF UNOBSERVABLES: A COMBINED REVEALED AND STATED PREFERENCE APPROACH[^]

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IDENTIFYING DEMAND PARAMETERS IN THE PRESENCE OF UNOBSERVABLES: A COMBINED REVEALED AND STATED PREFERENCE APPROACH

Abstract: We develop a combined revealed and stated preference approach to identify discrete choice demand parameters in the presence of unobserved determinants of choice. Our approach overcomes difficulties associated with small choice sets, multicollinearity, and endogenity that arise with revealed preference only approaches. To illustrate our framework, we revisit two Canadian moose hunting data sets. Our empirical results suggest the potential gains of fusing revealed and stated preference data, but they also suggest its limitations when the data generating processes for the data sources differ.

Key words: discrete choice models, non-market valuation, selection

JEL classifications: Q51, Q26, C25

I) Introduction

Estimating the demand for quality-differentiated goods such as recreation sites or residential location is a common problem in environmental economics. Typically, the analyst is interested in understanding the role of environmental quality in individual choice, and assessing how well-being changes when exogenous attributes change due to policy or other interventions. This type of inference problem requires that the analyst estimate a structural model of behavior, and discrete choice random utility maximization (RUM) models have become the standard approach in much of this literature.

Within this context, Timmins and Murdock (forthcoming) and Bayer, Keohane, and Timmins (2006) recently implemented a discrete choice econometric strategy that exploits revealed preference (RP) data to identify the behavioral and welfare effects of environmental quality while controlling for unobserved determinants of choice. Their RP approach is particularly appealing when significant preference heterogeneity exists and observed commodity attributes are endogenous, i.e., correlated with unobserved attributes. Three limitations with their strategy are that it requires: 1) data sets with many commodities (sites in the recreation context, houses or neighborhoods in the hedonic context); 2) variation in observed attributes across the objects of choice; and 3) instruments for endogenous variables. In this paper, we develop an alternative econometric strategy that exploits both RP and stated preference (SP) data to identify environmental quality's effect on behavior and welfare while controlling for unobserved attributes. The advantages of our proposed strategy are that it can be used in applications with small choice sets (14 and 11 in our empirical applications), exploits the experimental design embedded in the SP data to overcome multicollinearity problems often present in RP data, and does not require instruments for endogenous variables. Like Murdock

(2006), our strategy also controls for observable and unobservable preference heterogeneity. To our knowledge, our empirical models represent the most detailed accounting of both preference and attribute heterogeneity in combined RP/SP environmental applications.

The notion of "fusing" revealed and stated preference data has gained broad acceptance in environmental economics, transportation, and marketing over the past decade (see Whitehead et al. 2005 for a recent review). Combining RP decisions made at current conditions with SP choices made under hypothetical conditions within a unified behavioral model can aid in identifying the role of observable attributes in preferences. In the discrete choice context, the SP data is often generated with choice experiments (Kanninen, 2007). Choice experiments present individuals a series of hypothetical alternatives with randomly assigned attributes and ask them to state their preferred alternatives. Past experience with combining RP and choice experiment SP data suggests that combined RP/SP models can identify structural parameters associated with observed attributes that are not identified with just RP data (Adamowicz et al., 1994, 1997). For credible identification, however, it is critical that the analyst control for unobserved attributes associated with the RP objects of choice as well as observed and unobserved preference heterogeneity. To varying degrees, these issues have been dealt with in previous RP/SP studies. Our econometric strategy, however, addresses them in a more comprehensive and systematic manner.

To illustrate our proposed RP/SP econometric model, we reconsider two empirical applications related to recreational moose hunting in Canada (Adamowicz et al., 1997; Haener et al., 2001). The RP data consists of seasonal trip data to 14 wildlife management units in Alberta and 11 wildlife management zones in Saskatchewan. The SP data was constructed by presenting the same sample of individuals in the RP data a series of choice experiments related to hypothetical moose hunting sites with exogenously varying attributes. Our empirical results suggest the potential gains from fusing these different data sources in terms of parameter identification. They also suggest how the inclusion of controls for unobserved site attributes as well as observed and unobserved preference heterogeneity can generate improvements in statistical fit. Our richer models of heterogeneity come with a cost, however; tests for parameter consistency across the RP and SP data are routinely and strongly rejected. This finding runs counter to recent experimental evidence (List et al., 2006; Taylor et al., 2006). It also contradicts previous published findings that employ the same moose hunting data we use with more parsimonious specifications. Our findings raise important questions for how welfare evaluations should be conducted in these instances, and we consider a menu of plausible approaches to construct welfare measures that future researchers might find useful.

The remainder of the paper is organized as follows. In the next section we discuss identification issues arising with RP, SP, and combined RP/SP data. Section III describes the econometric specification which we apply to the Canadian moose hunting data sets described in Section IV. We then present our results, and close with a discussion of the practical implications of our findings for future research.

II) Identification with RP and SP data

In this section we develop a general discrete choice econometric model and discuss identification issues arising with RP data, SP data, and combined RP/SP data. We adopt notation that closely follows Berry, Levinsohn, and Pakes (2004) who first raised the identification issues we are concerned with in the context of RP applications. For concreteness, we couch our model in the context of travel cost models of recreation site choice (Phaneuf and Smith, 2005), although the model applies in principle to many different choice settings (e.g., housing, automobiles).

In the recreation site choice context, each individual is assumed to choose one of J sites as a recreation destination on a given trip. Recreation sites are differentiated in terms of their attributes and costs of visitation (travel costs) for each individual. To model such choices within a random utility maximization (RUM) framework, we begin by specifying the individual's conditional indirect utility function:

$$V_{ijt} = \boldsymbol{x}_{ijt}^{\mathsf{T}}\tilde{\boldsymbol{\gamma}}_i + \boldsymbol{x}_j^{\mathsf{T}}\tilde{\boldsymbol{\beta}}_i + \xi_j + \mu\varepsilon_{ijt}, \tag{1}$$

where i indexes individuals, j indexes the J sites, and t indexes trip occasions for individual i. The above specification assumes that utility is additive in four components: 1) $\boldsymbol{x}_{ijt}^{\mathsf{T}}\tilde{\boldsymbol{\gamma}}_i$, an index of observed site attributes, $\boldsymbol{x}_{ijt}$, that vary across individuals, trips, or both individuals and trips (e.g., travel costs, perceived or time-varying environmental quality measures); 2) $\boldsymbol{x}_j^{\mathsf{T}}\tilde{\boldsymbol{\beta}}_i$, an index of observed attributes, $\boldsymbol{x}_j$, that vary only across sites (e.g., time-invariant objective measures of environmental quality, availability of bath rooms or boat ramps); 3) ξ_j, an alternative specific constant (ASC) that controls for unobserved attributes that vary across sites and may be correlated with observed attributes; and 4) $\mu\varepsilon_{ijt}$, the product of an unobserved, idiosyncratic normalized error that varies across individuals, sites, and trips and a scale parameter μ. Although we allow for possible correlations among the first three components, we make the identifying assumption that ε_{ijt} is orthogonal to all other components in V_{ijt}. Furthermore, we specify the parameters in the $\boldsymbol{x}_{ijt}^{\mathsf{T}}\tilde{\boldsymbol{\gamma}}_i$ and $\boldsymbol{x}_j^{\mathsf{T}}\tilde{\boldsymbol{\beta}}_i$ indexes, $\tilde{\boldsymbol{\gamma}}_i$ and $\tilde{\boldsymbol{\beta}}_i$, to vary systematically and randomly across individuals:

$$\tilde{\gamma}_i = \overline{\gamma} + \mathbf{z}_i^{\mathsf{T}}\gamma^0 \;\; + \mathbf{v}_i \cdot \gamma^v$$
$$\tilde{\boldsymbol{\beta}}_i = \overline{\boldsymbol{\beta}} + \mathbf{z}_i^{\mathsf{T}}\boldsymbol{\beta}^0 + \boldsymbol{u}_i \cdot \boldsymbol{\beta}^u,$$

where ($\overline{\gamma}, \overline{\boldsymbol{\beta}}$) are parameter vectors of main or average effects, $\boldsymbol{z}_i$ is a vector of demographics, ($\gamma^0, \boldsymbol{\beta}^0$) are parameter matrices of interaction effects, ($\boldsymbol{v}_i, \boldsymbol{u}_i$) are normalized random effects that are independent of $\mathbf{z}_i$, ($\gamma^v, \boldsymbol{\beta}^u$) are vectors of standard errors for the random effects, and '$\cdot$' denotes scalar multiplication. The structure of preferences in (1) is very general in that it allows for systematic and random variation in preferences as well as unobserved site characteristics that influence choice. For later discussion, it is useful to collapse all elements in (1) that are common to site j and equal across individuals and time into a scalar δ_j:

$$\delta_j = \boldsymbol{x}_j^{\mathsf{T}}\overline{\boldsymbol{\beta}} + \xi_j, \tag{2}$$

and to group the remaining terms in the following way:

$$\boldsymbol{X}_{ijt} = \begin{bmatrix} \boldsymbol{x}_{ijt}^{\mathsf{T}} & \boldsymbol{x}_{ijt}^{\mathsf{T}} \cdot \boldsymbol{z}_i(1,1) & \boldsymbol{x}_{ijt}^{\mathsf{T}} \cdot \boldsymbol{z}_i(2,1) & \cdots & \boldsymbol{x}_j^{\mathsf{T}} \cdot \boldsymbol{z}_i(1,1) & \boldsymbol{x}_j^{\mathsf{T}} \cdot \boldsymbol{z}_i(2,1) & \cdots \end{bmatrix}^{\mathsf{T}}$$
$$\boldsymbol{\beta} = \begin{bmatrix} \overline{\gamma}^{\mathsf{T}} & \gamma^0(.,1)^{\mathsf{T}} & \gamma^0(.,2)^{\mathsf{T}} & \cdots & \boldsymbol{\beta}^0(.,1)^{\mathsf{T}} & \boldsymbol{\beta}^0(.,2)^{\mathsf{T}} & \cdots \end{bmatrix}^{\mathsf{T}}$$

$$\begin{aligned} \boldsymbol{e}_i &= [\boldsymbol{v}_i^{\mathsf{T}} \quad \boldsymbol{u}_i^{\mathsf{T}}]^{\mathsf{T}} \\ \boldsymbol{\sigma} &= [(\gamma^v)^{\mathsf{T}} \quad (\boldsymbol{\beta}^u)^{\mathsf{T}}]^{\mathsf{T}} \\ \boldsymbol{W}_{ijt} &= [\boldsymbol{x}_{ijt}^{\mathsf{T}} \quad \boldsymbol{x}_j^{\mathsf{T}}]^{\mathsf{T}}, \end{aligned} \tag{3}$$

so that the conditional utility function can be written compactly as

$$V_{ijt} = \boldsymbol{X}_{ijt}^{\mathsf{T}}\boldsymbol{\beta} + \boldsymbol{W}_{ijt}^{\mathsf{T}}(\boldsymbol{e}_i \cdot \boldsymbol{\sigma}) + \delta_j + \mu\varepsilon_{ijt}. \tag{4}$$

We now discuss how the analyst can identify the structural parameters in (1) with RP, SP, and combined RP and SP data. Consider first the case where only RP data is available. Building on an econometric strategy developed by Berry, Levinsohn, and Pakes (2004), Timmins and Murdock (forthcoming), Bayer, Keohane, and Timmins (2006), and Murdock (2006) propose the following two-step approach. In the first step, the analyst estimates the parameters in (4) through maximum likelihood. This requires normalizing the scale parameter μ to one and one of the ASCs (say δ_1) to zero with no loss in generality. If several random effects are included in the empirical specification, simulation-based estimation will be required. Also, in applications with many sites, estimating the full set of ASCs may require the use of a contraction mapping algorithm developed by Berry (1994).

If the analyst is concerned with policies involving changes in $\boldsymbol{x}_{ijt}$ such as site loss scenarios, the first-stage estimates will have recovered enough parameters for welfare analysis.[1] However, if the analyst's concern centers on exogenous policy shocks involving changes to observed attributes that only vary across sites, a second-stage regression based on the specification in (2) above will be necessary. In these instances, the estimated ASCs are regressed on the observed attributes $\boldsymbol{x}_j$. Implementation of this regression may be confounded by a number of factors that are of practical importance in environmental applications. First, for the regression to satisfy the rank condition, no observed attributes can be linear combinations of the other attributes, and the number of ASCs (i.e., the number of sites minus one) must be greater than the dimension of $\boldsymbol{x}_j$. Second, past experience suggests that the number of ASCs must be substantially greater than the dimension of $\boldsymbol{x}_j$ for parameters to be precisely estimated.[2] Third, if there are correlations between the unobserved site attributes ξ_j and $\boldsymbol{x}_j$, instruments will be required. Timmins and Murdock and Bayer, Keohane, and Timmins, building on earlier work by Bayer and Timmins (forthcoming), discuss how the structure of the discrete choice model can be used to develop instruments for observed attributes that are determined through social interactions such as congestion. Otherwise, the analyst will need to find other sources for variables that are correlated with the observables but uncorrelated with the unobservables. The use of instruments in the second stage, particularly if they have little identifying power, also exacerbates the need for a large choice set to precisely estimate the structural parameters. For all of these reasons, identification may be elusive in environmental applications that exploit only RP data.

SP methods can in principle be used to identify parameters in (1). For example, choice experiments (Kanninen, 2007) can be designed that present respondents with a series of hypothetical recreation sites that vary exogenously in the level of measurable attributes. For each hypothetical choice set, respondents are asked to state which if any of the sites they would choose to visit. Recent experimental evidence (Taylor et al., 2006; List et al., 2006) suggests that the process by which experienced individuals make these hypothetical choices is very similar to how they make similar choices in consequential (i.e., real) situations except for a tendency to choose the 'opt-out' or 'no trip' alternative more frequently. These findings suggest that SP data can be used to identify at least some of the relevant structural parameters in (1).

Identification with SP data is achieved as follows. Similar to the RP choice context, the analyst specifies a conditional indirect utility function for each hypothetical site j and choice question c of the form:

$$
\begin{aligned}
V_{ijc} &= \boldsymbol{x}_{ijc}^{\mathsf{T}}(\tilde{\boldsymbol{\gamma}}_i + \tilde{\boldsymbol{\beta}}_i) + \mu^{*}\varepsilon_{ijc} \\
\tilde{\boldsymbol{\gamma}}_i &= \overline{\boldsymbol{\gamma}} + \mathbf{z}_i^{\mathsf{T}}\boldsymbol{\gamma}^0 + \mathbf{v}_i \cdot \boldsymbol{\gamma}^v \\
\tilde{\boldsymbol{\beta}}_i &= \overline{\boldsymbol{\beta}} + \mathbf{z}_i^{\mathsf{T}}\boldsymbol{\beta}^0 + \boldsymbol{u}_i \cdot \boldsymbol{\beta}^u.
\end{aligned}
\tag{5}
$$

There are three significant differences between the SP and RP preference specifications in (5) and (1). First, due to the random assignment of all observed attributes across individuals, alternatives, and choice experiments, the observed attributes that fall in the $\boldsymbol{x}_j$ vector in the RP data instead fall in the $\boldsymbol{x}_{ijc}$ vector in the SP data. This implies that *all* main and interaction effects in $\tilde{\boldsymbol{\gamma}}_i$ and $\tilde{\boldsymbol{\beta}}_i$ can be identified without relying on a second-stage regression. Second, the scale parameter for the idiosyncratic error, μ^*, may differ between the RP and SP data sets (Swait and

Louviere, 1993). This possibility has little practical import when using just SP data to identify the structural parameters in (1) because arbitrary rescaling of all parameters has no effect on welfare analysis. It does, however, have important implications for pooling RP and SP data as we discuss below. Third, whereas the RP objects of choice consist of observed and unobserved attributes that may be correlated, the SP objects of choice consist only of observed attributes. The absence of a role for unobserved site attributes reflects that fact that environmental choice experiment applications do not "brand" each choice alternative as a real (but experimentally altered) recreation site with which respondents have experience.[3] This implies that the ξ_j's in (1) cannot be identified with the SP data, and thus the full structure of preferences cannot be recovered from SP data alone. As a result, welfare analysis is confounded unless the analyst restrictively assumes that unobserved attributes have no impact on choice.

Jointly estimating preferences with RP and SP data, however, permits identification of all structural parameters without relying on a second-stage regression or restrictions on the unobserved attributes' parameter values. To see this, recall that the identification difficulties with RP only approaches arise because both the unobserved attributes and the main effects associated with observed attributes that only vary across sites are captured in the first-stage estimated ASCs. In other words, the linear combination $\boldsymbol{x}_j^{\mathsf{T}}\bar{\boldsymbol{\beta}} + \xi_j$ is identified in the first stage of the RP only estimation strategy, not the $\bar{\boldsymbol{\beta}}$ and ξ_j parameters. Recall that with SP data, all parameters in $\tilde{\boldsymbol{\gamma}}_i$ and $\tilde{\boldsymbol{\beta}}_i$ (including $\bar{\boldsymbol{\beta}}$) are identified off of the exogenous variation in all observed attributes embedded in SP experimental design. However, if the RP data identifies $\boldsymbol{x}_j^{\mathsf{T}}\bar{\boldsymbol{\beta}} + \xi_j$ and the SP data identifies $\bar{\boldsymbol{\beta}}$, the combination of RP and SP data sets jointly identify $\bar{\boldsymbol{\beta}}$ and ξ_j. This implies that welfare analysis is feasible for scenarios involving changes in all site attributes in applications with small choice sets, little or no variation in observed attributes, and unavailable instruments for the endogenous observed attributes.

It is worth emphasizing that, in addition to the experimental design embedded in the SP choice experiments, identification with combined RP/SP data depends critically on the assumption that a common data generating process gives rise to both RP and SP choices. The cross-equation restrictions implied by the common data generating process can be relaxed in some ways – differences in RP and SP scale for the idiosyncratic error term can be introduced, and the frequency of the 'opt-out' alternative being chosen in the SP choice experiments can be controlled. The parameters identified by both data sources (i.e., the main effects for site attributes that vary across individuals or time as well as all interaction and random effects) must be restricted to be equal for the approach to have identifying power. As stated above, some experimental evidence suggests that these parameters may converge across RP and SP data sources, especially with experienced respondents. The evidence from past RP/SP studies, however, is mixed (Adamowicz et al., 1994; Earnhart, 2001). From a statistical perspective, parameter convergence across RP and SP sources is a testable hypothesis, and we therefore conduct likelihood ratio tests to evaluate these cross-equation restrictions in our empirical applications.

III) Econometric specification

To transform the behavioral model outlined in the previous section into an estimable econometric model, we make the following distributional assumptions. First, we assume the idiosyncratic errors in (1) and (5) are independent and identically distributed type I extreme value draws. This assumption, omnipresent in the discrete choice literature, implies that the RP and SP conditional choice probabilities have the convenient logit form. We allow for differences in scale between the RP and SP data by incorporating the parameter ratio μ/μ^* in the SP conditional probabilities (Louviere and Swait, 1993).[4] Second, we assume the individual-specific random effects in the parameter vectors, ($\boldsymbol{v}_{i,}\, \boldsymbol{u}_i$) are independent and identically distributed standard normal. To move from the conditional choice probabilities implied by the type I extreme value assumption to the unconditional probabilities used in estimation, we employ simulation (Train, 2003). We then use the method of maximum likelihood to estimate the structural parameters. Our estimation code is available upon request.

Our empirical applications involve choice sets that are relatively small (14 and 11 sites, respectively), and thus estimating a full set of ASCs along with the main, interaction, and scale ratio parameters is relatively straightforward using standard gradient-based search techniques. For large choice set applications, employing Berry's (1994) contraction mapping to numerically solve for the implied ASCs may be necessary.

IV) Data

To illustrate the logic of our approach and demonstrate its potential value, we reconsider two related data sets previously analyzed by Adamowicz et al. (1997) and Haener et al. (2001). The first examines preferences among moose hunters in the Canadian province of Alberta for visits to 14 wildlife management units (WMUs) and their associated site attributes. This data set provides information on the actual visits made by 271 moose hunters to the WMUs as well as answers to a series of choice experiment questions that solicited hypothetical choices among two generic hunting sites and an 'opt-out' option. The experimental design included attributes based on distance from home, road quality on which the person traveled to reach the site, access conditions, encounters with other hunters, forestry activity at the site, and the local moose population. Expected values for these variables at the 14 WMUs are also available corresponding to the time that the observed visitation behavior was recorded. The top half of table 1 provides a summary of the variables and their RP means that are available in both the RP and SP data as well as information that is available on the individual survey respondents.

The second data set is similar to the first but examines preferences among Saskatchewan moose hunters for visits to 11 wildlife management zones (WMZs). 532 hunters provided information on the trips they made to the 11 sites as well as answers to a set of choice experiment questions that were similar to the Alberta SP application. Attributes included distance, access conditions to the site, encounters with other hunters, forestry activity, the local moose population, and variables describing other species present at the site. The bottom half of table 1 provides a description and summary of the variables as defined and used in the analysis.[5]

V) Results

The results of our analysis using the Alberta data are presented in tables 2 and 3 and for the Saskatchewan data in tables 4 and 5. Table 2 reports parameter estimates with the Alberta data set for several systematically and randomly varying parameter specifications. Each row of the table corresponds to a separate variable that we wish to identify. These include a travel cost term, attributes of the 14 recreation sites, interactions between three individual characteristics and each site attribute, a random coefficient for each attribute, a scale ratio parameter, and ASCs for each recreation site less one for normalization. We present parameter estimates and t-statistics for five models arrayed across the ten columns. From left to right, these include a RP only model without ASCs, a RP only model with ASCs, a SP only model, a combined RP/SP model without ASCs, and our most general model that uses both RP and SP data with ASCs.

In general, we find parameter signs and significance that are intuitive and match previous analyses using this data. For all five models, both the inclusion of interaction terms and random effects in the parameter specifications jointly add considerable explanatory power (see the reported p-values at the bottom of table 2), although many of the interaction terms are not individually statistically significant.[6] Focusing first on the RP data only model without ASCs in columns one and two, we find that a number of parameters (*No Trail*, *4WD Trail*, *On Foot*, *On ATV*) are not identified due to multicollinearity in the RP data. The addition of ASCs in columns three and four results in a more severe identification problem as the main effects associated with all observed site attributes except *Unpaved* (which varies across sites and individuals) are confounded with the unobserved site attributes. Moreover, a second-stage OLS regression of the ASCs on the observed attributes did not produce significant or plausibly signed estimates.[7] As a result, welfare analyses involving changes in any attribute besides *Travel Cost* and *Unpaved* are not possible. This finding might lead one to prefer the more parsimonious specification without ASCs, but a likelihood ratio test strongly suggests that the gains in statistical fit arising from the addition of the ASCs are significant (p-value <0.0001).

Columns five and six report parameter estimates for the SP data only model. For this specification, an *Outside Dummy* variable is interacted with demographics and a random effect to account for the inclusion of the ‘opt-out’ option. Confirming our claim in section II, the experimental design embedded in the choice experiments permits identification of all main, interaction, and random effect parameters for the observed attributes. However, the SP data does not identify the role of unobserved attributes that are specific to the WMUs. Unless the analyst restrictively assumes that unobserved attributes have no influence on moose hunters' decisions, welfare analysis is not feasible with the SP only specification.

Finally, columns seven through ten report estimates for the combined RP/SP models. Fusing the data sources facilitates identification of all main, interaction, and random effects for the observed attributes. Comparing the combined RP/SP models with and without ASCs, we find a substantial and statistically significant improvement in fit generated by the addition of ASCs (p-value < 0.0001). Combined with the results in columns three and four, these results highlight the significant gains in statistical fit arising from controlling for unobserved site attributes in the RP data. Moreover, although the signs of the main effects parameters generally do not change with the addition of ASCs, their magnitudes change substantially (see, for example, the *Old Trail*, *<1 Moose,* and *3/4 Moose* variables). This finding suggests that correlations between observed and unobserved attributes are present in the data and lead to biased estimates.

Based on the RP only, SP only, and combined RP/SP results, it is possible to test whether the cross-equation restrictions embedded in the combined RP/SP model are statistically valid as is common in the RP/SP literature. A likelihood ratio test suggests that we can strongly reject the cross-equation restrictions (p-value < 0.0001). This empirical finding contradicts Adamowicz et al.'s (1997) earlier findings with the same Alberta data that use more parsimonious models.[8] It also runs counter to the experimental evidence from Taylor et al. (2006) and List et al. (2006) that find pooling the observed attribute parameters across the RP and SP data cannot be statistically rejected. Upon discovery of this empirical result, we first conjectured that it may reflect that our rich models of observed and unobserved heterogeneity are over-fitting the RP and SP data and spuriously leading us to reject the cross-equation restrictions.

We therefore considered a number of more parsimonious specifications where we dropped interaction terms selectively and in total (see the bottom of table 2). P-values from these specifications similarly implied that the cross-equation restrictions could be strongly rejected. We also conjectured that measurement error in the RP data's observed attributes might explain our rejection of the cross-equation restrictions. Recall that the RP attributes are averages across relatively large land areas, often several thousand hectares in size. It is difficult for us to statistically evaluate this conjecture, although we suspect that it plays some role in our findings. Also, we recognized that one could logically conclude from our results that our maintained assumption that the RP and SP data generating processes are the same is invalid, perhaps due to hypothetical or strategic bias in the SP data.[9] To the degree that differences in the behavioral models that generated the RP and SP data exist, they represent significant limitations for our identification strategy.

To further examine the differences between models, table 3 reports partial-equilibrium[10] welfare estimates for three scenarios:

Scenario #1: Loss of Site – remove WMU #344 from the choice set (site #5).

Scenario #2: Reduction in Moose Population – move from > 4 moose per day to 3-4 moose per day at WMU #348 (site #7).

Scenario #3: Increase in Moose Population – move from < 1 moose per day to 1-2 moose per day at WMU #344 (site #5).

Following von Haefen (2003), each estimate is constructed conditionally on the observed choices. For the specification using only RP data and no ASCs in row one, we find that all welfare estimates are identified but have large standard errors. This contrasts with the RP data with ASCs models in row two, which are only identified for the site loss scenario because the moose attributes' main effects are confounded with unobserved site attributes. The SP data only estimates in row three are qualitatively similar and far more precise relative to their RP only counterparts. They rely, however, on the restrictive and unrealistic assumption that unobserved determinants of sites do not influence choice.

Turning to the combined RP/SP models, rows four and five of table 3 report estimates based on the specifications with cross-equation restrictions (columns seven through ten in table 2). For the site loss and reduced moose population scenarios, we find qualitatively similar estimates. The combined RP/SP reduced moose population scenario point estimates are smaller than those exploiting either RP or SP data only, but the latter have significantly larger standard errors. The increased moose population scenarios are qualitatively different with and without ASCs. The inclusion of ASCs results in much a larger point estimate ($61.02 versus $2.99), but the larger estimate also has a substantially larger standard error (20.6 versus 2.37). This empirical finding is consistent with Murdock's (2006) Monte Carlo results that the exclusion of ASCs can lead to biased welfare estimates and spurious precision.

Because we strongly rejected the hypothesis that the cross-equation restrictions embedded in the combined RP/SP models hold, we also considered some alternative strategies for constructing welfare measures that relied on both data sources but did not impose cross-equation restrictions. The first strategy used the RP data only estimates when available, and used the SP data only estimates to "fill in" unidentified estimates. This was accomplished by treating the ratio of travel costs parameters in the RP only and SP only models as an estimate of the scale parameter ratio, and using this estimated ratio to rescale the SP estimates and arrive at fill-in values for the missing RP only estimates. The second strategy follows Swait, Louviere, and Williams (1994) and combines the SP only estimates with estimates of the ASCs constructed with the RP data. These ASC estimates were constructed by "concentrated" maximum likelihood, i.e., fixing the main, interaction, and random effects at their estimated SP only values and conditionally estimating a full set of ASCs with the RP data. Welfare results based on these two strategies are reported in the additional results section of table 3. We find that these estimates are qualitatively similar for the site loss and reduced moose population scenarios, but smaller and more precisely estimated for the increased moose population scenario. The divergence in the increased moose population estimates reflects the fact that the key parameters (*<1 Moose* and *1/2 Moose*) are qualitatively different across the RP only, SP only, and combined RP/SP models. We do not take a stand on which of these estimates is most defensible, but instead view the range of estimates as providing welfare bounds for this scenario.[11]

Tables 4 and 5 report a similar set of findings with the Saskatchewan data which we briefly summarize here. Across the five different models reported in tables 4 and 5, we find the inclusion of interaction and random effects to be strongly statistically significant. The parameter estimates reported in columns one through six in table 4 also suggest that several parameters are not identified when either RP or SP data is used alone. Fusing the RP and SP data generates similar gains in terms of parameter identification (see columns seven through ten), and the inclusion of a full set of ASCs generates a significant improvement in fit as well as qualitatively different welfare estimates (e.g., the moose main effects). Likelihood ratio tests of the cross-equation restrictions embedded in the combined RP/SP models are again strongly rejected.

Table 5 reports welfare estimates for three policy scenarios:

<u>Scenario #1: Loss of Site</u> – remove WMZ #59 from the choice set.

<u>Scenario #2: Reduction in Moose Population</u> – move from > 3 moose per two days to 1 moose per day at WMZ #59.

Scenario #3: Increase in Moose Population – move from 1 moose per two days to 1 moose per day at WMZ #66.

The pattern of these estimates is similar to what we found with the Alberta data. Only the site loss welfare estimates could be constructed with the RP data due to multicollinearity in the moose attributes, and the SP only welfare estimates assume that unobserved attributes have no role in determining choice. Some differences between the combined RP/SP welfare measures arise for specifications with and without ASCs, especially in terms of the increased moose population scenario. Comparisons across welfare estimates for combined RP/SP models that impose cross-equation restrictions and those that do not suggest a range of values for all three scenarios, although the estimates based on the RP models with SP fill-ins are notably imprecise.

VI) Conclusion

We draw four general conclusions from our conceptual and empirical findings in this paper. First, accounting for heterogeneity – whether it arises from unobserved site attributes or observed and unobserved preference variation – is an issue of substantial importance in non-market valuation applications. Results from our moose hunting applications as well as previous results by Murdock (2006), Timmins and Murdock (forthcoming), and Bayer, Keohane, and Timmins (2006) strongly suggest that accounting for heterogeneity can improve statistical fit, reduce bias, and alter policy implications. Second, our combined RP/SP approach to identifying preference parameters in the presence of unobserved determinants of choice represents a feasible and in many ways attractive alternative to RP only approaches. By fusing these data sources, we circumvented the limitations associated with RP only two-step estimators that require large choice sets, variation in the observed attributes, and instruments for endogenous attributes.

Third, our combined RP/SP approach has limitations too. Most importantly, it requires additional SP data that, with few exceptions, must be generated by the same behavioral process that generated the RP data. Finally, when the assumption of a common data generating process across the RP and SP data sources is rejected, the appropriate strategy for constructing welfare measures is unclear. For roughly half of the policy scenarios we considered, we found qualitatively similar estimates across three arbitrary but plausible approaches to welfare construction. For the other scenarios, our three approaches to constructing welfare measures implied bounds that may be sufficiently informative for policy.

Finally, it is our view that both the RP only and combined RP/SP approaches to identifying preference parameters in the presence of unobserved attributes have strengths and weaknesses. We believe that neither approach is strictly preferred to the other, and data environments will likely dictate which approach is adopted in future applications. Having said this, we also believe that future research should empirically compare both approaches with a common data set to ascertain whether they generate qualitatively similar parameter and policy inferences. Such a comparison would demand a very rich data set – one with SP data as well as RP data with many objects of choice, sufficient orthogonality in attribute space, and instruments for endogenous attributes. It would, however, have the potential to clarify the performance of both approaches in a relatively controlled data environment.

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Table 1: Variable Definitions

1993 Alberta Data (14 Sites, 271 Observations)

Variable	Definition	RP Mean
Site Characteristics		
Travel Cost[1]	Round trip travel cost	$219.71
Road quality (excluded category – mostly paved, some gravel or dirt roads)		
Unpaved	Some paved, mostly gravel or dirt roads (effects coded)	-0.82
Hunter access (excluded category – newer trails, passable with two-wheel-drive vehicle)		
No Trail	No trails, cutlines, or seismic lines (effects coded)	-0.21
Old Trail	Old trails, passable with ATV (effects coded)	0.21
4WD Trail	Newer trails, passable with four-wheel-drive vehicle	0.14
Hunter congestion (excluded category – encounters with other hunters on trucks)		
No Hunters	Encounters with no other hunters (effects coded)	-0.64
On Foot	Encounters with other hunters on foot (effects coded)	-0.64
On ATV	Encounters with other hunters on ATVs (effects coded)	-0.29
Forestry activity (excluded category – some evidence of recent forestry activity)		
No Logging	No evidence of recent forestry activity (effects coded)	0.57
Moose variables (excluded category – evidence of 4 or more moose per day)		
< 1 Moose	Evidence of < 1 moose per day (effects coded)	0.14
1/2 Moose	Evidence of 1 to 2 moose per day (effects coded)	0.50
3/4 Moose	Evidence of 3 to 4 moose per day (effects coded)	0.07
Individual Characteristics		Sample Mean
Total Trips	Average number of trips to all sites	3.62
Gen Hunt Exp	Years of hunting experience (count)	20.2
Edmonton	Dummy for Edmonton resident	0.45
HS Diploma	Dummy for high school diploma	0.91

1994 Saskatchewan Data (11 Sites, 532 Observations)

Variable	Definition	RP Mean
Site Characteristics		
Travel Cost[1]	Round trip travel cost	$251.49
Hunter access (excluded category – access on foot or ATV)		
2WD Access	passable with two-wheel-drive vehicle (effects coded)	0.55
4WD Access	passable with four-wheel-drive vehicle (effects coded)	0.45
Hunter congestion (excluded category – encounters with other hunters on ATVs)		
No Hunters	Encounters with no other hunters (effects coded)	-0.45
On Foot	Encounters with other hunters on foot (effects coded)	0.09
Forestry activity (excluded category – some evidence of recent forestry activity)		
No Logging	Little or no evidence of recent forestry activity (effects coded)	0.27
Moose variables (excluded category – evidence of 3 moose every 2 days)		
< 1 Moose	Evidence of < 1 moose every 2 days (effects coded)	0.55
1 Moose	Evidence of 1 moose per day (effects coded)	0.18
Wildlife species (excluded category – sightings of common wildlife, 1-2 previously unseen species, and possibly rare or endangered species)		
Common Species	Only sightings of common wildlife (effects coded)	-0.09
Unseen Species	Sightings of common wildlife and 1-2 previously unseen species (effects coded)	-0.55
Individual Characteristics		Sample Mean
Total Trips	Average number of trips to all sites	1.37
Gen Hunt Exp	Years of hunting experience (count)	23.23
Urban	Dummy for residents of Whitecourt, Hinton, Edson, or Drayton Valley	0.81
HS Diploma	Dummy for high school diploma	0.89

[1] For the Alberta SP data, travel costs were generated assuming travel distances of 50, 150, 250, and 350 kilometers. For the Saskatchewan SP data, travel costs were generated assuming travel distances of 75, 250, and 425 kilometers.

Table 2 – Alberta Parameter Estimates

	RP Data Only, No ASCs		***RP Data Only, ASCs***		***SP Data Only***		***RP & SP Data, No ASCs***		***RP & SP Data, ASCs***	
Log-Likelihood	-1,465.7		-1,386.4		-3,010.0		-4,817.8		-4,521.6	
	Est.	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***
RP/SP Scale Ratio	-	-	-	-	-	-	**0.448**	5.73	**0.240**	7.31
Travel cost	-0.615	-1.73	**-1.305**	-4.48	**-0.701**	-12.3	**-1.148**	-5.45	**-1.935**	-7.69
Unpaved	1.202	1.46	-0.200	-0.23	0.001	0.02	0.165	0.80	0.538	0.83
x Gen Hunt Exp	-0.019	-0.68	0.030	0.82	-0.005	-1.29	-1.035	-1.30	-2.164	-1.78
x Edmonton	-	-	-	-	**0.169**	2.34	0.274	1.78	**0.595**	2.17
x HS Diploma	**-3.026**	-2.15	**-3.355**	-2.26	0.008	0.06	-0.311	-0.93	-0.353	-0.60
x Random Effect	**1.898**	3.25	**1.974**	6.51	0.109	1.01	**0.356**	2.77	**0.554**	2.19
No Trail	-	-	-	-	-0.105	-0.47	-0.444	-1.01	-1.936	-1.48
x Gen Hunt Exp	-	-	-	-	0.001	0.11	0.927	0.51	0.622	0.19
x Edmonton	-	-	-	-	**-0.763**	-4.25	**-1.366**	-3.23	**-2.219**	-3.32
x HS Diploma	-	-	-	-	0.149	0.45	0.015	0.01	1.679	1.46
x Random Effect	-	-	-	-	**0.820**	8.18	**1.039**	5.22	**1.861**	6.12
Old Trail	0.758	0.80	-	-	0.313	1.51	0.253	0.40	1.861	1.84
x Gen Hunt Exp	0.032	0.52	-0.023	-0.77	0.005	0.44	1.891	0.71	-1.079	-0.48
x Edmonton	0.813	0.79	**1.536**	3.58	0.322	1.68	**1.301**	2.76	**1.997**	2.92
x HS Diploma	1.828	0.77	1.525	0.73	-0.201	-0.51	0.634	0.98	-0.693	-0.93
x Random Effect	**2.041**	3.14	**2.202**	5.96	**0.919**	9.30	**1.364**	8.02	**1.863**	6.77
4WD Trail	-	-	-	-	0.150	0.87	-0.289	-0.49	0.737	0.98
x Gen Hunt Exp	-	-	-	-	0.001	0.14	1.012	0.40	2.158	1.04
x Edmonton	-	-	-	-	0.218	1.43	0.395	1.20	0.744	1.46
x HS Diploma	-	-	-	-	0.257	0.69	0.263	0.31	-0.658	-1.10
x Random Effect	-	-	-	-	**0.496**	6.23	**0.897**	3.93	**1.195**	4.61
No Hunters	**-2.485**	-3.75	-	-	**1.272**	7.93	**2.293**	5.08	**3.357**	2.81
x Gen Hunt Exp	0.001	0.07	-0.007	-0.28	**-0.017**	-2.85	**-2.750**	-2.08	**-4.725**	-2.05
x Edmonton	**0.971**	2.41	**2.685**	4.30	-0.027	-0.20	-0.002	-0.01	-0.232	-0.47
x HS Diploma	-0.765	-0.71	-0.452	-0.30	0.134	0.38	0.356	0.53	0.998	1.13
x Random Effect	**0.925**	2.34	**1.298**	5.93	**0.397**	4.62	**0.510**	3.50	**0.940**	2.64
On Foot	-	-	-	-	-0.127	-0.65	-0.180	-0.51	-1.216	-0.96
x Gen Hunt Exp	-	-	-	-	0.003	0.34	0.293	0.19	1.107	0.40
x Edmonton	-	-	-	-	**0.316**	2.11	0.524	1.73	0.909	1.77
x HS Diploma	-	-	-	-	-0.176	-0.50	-0.311	-0.48	0.750	0.69
x Random Effect	-	-	-	-	**0.382**	4.05	0.235	1.12	0.634	1.57
On ATV	-	-	-	-	**-0.430**	-2.27	**-1.541**	-3.92	-1.200	-1.16
x Gen Hunt Exp	-	-	-	-	0.003	0.42	1.309	0.88	1.296	0.62
x Edmonton	-	-	-	-	0.010	0.07	-0.022	-0.09	0.543	1.13
x HS Diploma	-	-	-	-	-0.223	-0.74	-0.460	-0.84	-0.591	-0.64
x Random Effect	-	-	-	-	0.251	1.87	**0.423**	2.66	**0.794**	2.88
No Logging	1.243	0.85	-	-	-0.049	-0.39	0.158	1.04	-0.242	-0.67
x Gen Hunt Exp	0.002	0.04	0.025	1.00	0.006	1.08	0.738	0.98	1.215	1.23
x Edmonton	-0.469	-0.36	-0.405	-0.79	-0.013	-0.15	0.206	1.50	**0.538**	2.43
x HS Diploma	-0.149	-0.09	0.202	0.11	-0.211	-1.59	-0.292	-1.07	-0.126	-0.47
x Random Effect	**1.735**	3.38	**1.083**	6.00	**0.321**	3.97	**0.379**	3.22	**0.632**	2.79

Boldface indicates statistical significance at the 5 percent level. All random coefficient estimates generated with 500 quasi-random draws.

Table 2 – Alberta Parameter Estimates (cont.)

	RP Data Only, No ASCs		*RP Data Only, ASCs*		*SP Data Only*		*RP & SP Data, No ASCs*		*RP & SP Data, ASCs*	
	Est.	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***
< 1 Moose	**-1.791**	-1.96	-	-	**-1.801**	-6.95	**-2.187**	-4.23	**-5.702**	-5.31
x Gen Hunt Exp	0.034	1.03	0.020	0.47	0.001	0.11	0.764	0.40	-1.795	-1.05
x Edmonton	0.091	0.14	0.075	0.08	-0.080	-0.41	**-0.948**	-2.21	-0.391	-0.91
x HS Diploma	**-6.227**	-3.53	**-7.284**	-4.81	-0.396	-0.69	-1.674	-1.14	-0.197	-0.35
x Random Effect	**2.130**	4.54	**3.253**	4.78	**0.773**	7.86	**1.913**	6.34	**1.669**	7.06
1/2 Moose	-0.485	-1.32	-	-	-0.101	-0.56	-0.002	-0.01	-0.990	-1.37
x Gen Hunt Exp	0.015	1.00	0.005	0.20	0.001	0.07	-2.036	-1.69	-1.205	-0.86
x Edmonton	**2.404**	5.70	**3.575**	5.21	-0.060	-0.42	**0.915**	3.26	**1.808**	5.30
x HS Diploma	0.215	0.24	0.727	0.59	-0.345	-0.97	0.642	1.03	0.088	0.15
x Random Effect	**1.558**	4.22	**1.661**	6.12	**0.517**	5.13	**0.890**	6.50	**1.721**	6.10
3/4 Moose	-1.405	-1.67	-	-	**0.591**	3.57	0.384	1.26	2.052	1.52
x Gen Hunt Exp	**0.068**	2.26	**0.065**	2.26	0.003	0.50	1.800	1.44	1.635	1.16
x Edmonton	0.003	0.00	0.133	0.15	0.001	0.00	-0.234	-0.92	-0.134	-0.42
x HS Diploma	1.950	1.89	2.085	1.90	0.193	0.61	0.724	1.46	-0.042	-0.03
x Random Effect	**2.071**	3.40	**2.935**	6.16	0.041	0.30	**0.711**	5.55	**1.196**	6.11
SP Outside Dummy	-	-	-	-	**-3.071**	-5.29	**-5.940**	-3.82	**-8.473**	-2.92
x Gen Hunt Exp	-	-	-	-	-0.014	-0.53	0.242	0.06	3.473	0.50
x Edmonton	-	-	-	-	-0.661	-1.47	-1.605	-1.49	-2.152	-1.27
x HS Diploma	-	-	-	-	0.494	0.77	**2.763**	2.16	**-3.330**	-1.98
x Random Effect	-	-	-	-	**2.518**	11.49	**4.533**	4.77	**9.200**	6.89
WMU #337 ASC	-	-	**2.443**	2.04	-	-	-	-	0.258	0.32
WMU #338 ASC	-	-	1.555	1.32	-	-	-	-	**-2.110**	-2.15
WMU #340 ASC	-	-	0.973	1.02	-	-	-	-	**1.508**	2.55
WMU #342 ASC	-	-	-1.526	-1.18	-	-	-	-	**5.417**	5.50
WMU #344 ASC	-	-	2.774	1.67	-	-	-	-	**7.149**	5.25
WMU #346 ASC	-	-	**3.653**	2.59	-	-	-	-	-0.916	-0.86
WMU #348 ASC	-	-	1.786	0.86	-	-	-	-	**-6.165**	-4.20
WMU #350 ASC	-	-	**3.816**	2.42	-	-	-	-	1.711	1.86
WMU #352 ASC	-	-	1.811	1.10	-	-	-	-	**5.609**	4.93
WMU #354 ASC	-	-	2.420	1.48	-	-	-	-	1.492	1.45
WMU #356 ASC	-	-	**4.598**	3.69	-	-	-	-	**4.317**	3.75
WMU #437 ASC	-	-	-2.784	-1.43	-	-	-	-	**-4.794**	-3.18
WMU #438 ASC	-	-	-2.907	-1.55	-	-	-	-	**-3.656**	-2.11
Heterogeneity Test P-Values										
H_0: Interact. = 0	< 0.001		< 0.0001		0.0062		< 0.0001		< 0.0001	
H_0: Random Eff. = 0	< 0.0001		< 0.0001		< 0.0001		< 0.0001		< 0.0001	
H_0: ASCs = 0	-		< 0.0001		-		-		< 0.0001	

RP/SP Pooling Test P-Values

Models w/	***Interact., Rand. Eff., ASCs***	***Interact., Rand. Eff.***	***Some Interact., Rand. Eff., ASCs***	***Interact., ASCs***	***Rand. Eff., ASCs***
H_0: Common RP & SP Param. Equal	< 0.0001	< 0.0001	< 0.0001	< 0.0001	< 0.0001

Boldface indicates statistical significance at the 5 percent level. All random coefficient estimates generated with 500 quasi-random draws. Alternative specific constant (ASC) for WMU #507 is excluded. For "Some Interact., Rand. Eff., ASCs" specification in RP/SP pooling test section, all interactions that were not statistically significant at $\alpha = 0.05$ in the most general model in columns 9-10 were set to zero.

Table 3 – Alberta Welfare Estimates (1993 Canadian Dollars)

Specification	***Loss of Site WMU # 344***		***Reduction in Moose Population at WMU #348 (> 4 moose/day to 3 to 4 moose/day)***		***Increase in Moose Population at WMU #344 (<1 moose/day to 1 to 2 moose/day)***	
	Mean	***Std. Err.***	***Mean***	***Std. Err.***	***Mean***	***Std. Err.***
RP Data Only – no ASCs	-$10.47	83.1	-$65.42	536	$8.82	123
RP Data Only – ASCs	-$4.89	1.40	-	-	-	-
SP Data Only	-$0.75	0.53	-$61.98	30.9	$2.88	1.40
RP/SP Data – no ASCs	-$5.36	0.94	-$22.04	3.91	$2.99	2.37
RP/SP Data – ASCs	-$4.18	0.47	-$17.01	1.88	$61.02	20.6
Additional Results:						
RP Data – ASCs with SP Data Fill-Ins	-$4.98	1.21	-$16.38	16.65	$1.10	3.00
SP Data – ASCs from RP Data	-$8.82	0.79	-$27.51	3.35	$19.86	7.36

Welfare measures are conditional on observed choice (von Haefen, 2003); based on 3,500 simulations with the first 500 discarded as burn-in.

Table 4 – Saskatchewan Parameter Estimates

	RP Data Only, No ASCs		***RP Data Only, ASCs***		***SP Data Only***		***RP & SP Data, No ASCs***		***RP & SP Data, ASCs***	
Log-Likelihood	-833.1		-828.1		-5,583.2		-6,658.2		-6,547.5	
	Est.	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***
RP/SP Scale Ratio	-	-	-	-	-	-	**0.310**	7.55	**0.115**	6.11
Travel cost	**-1.561**	-5.83	**-2.293**	-4.01	**-0.381**	-11.7	**-1.262**	-7.89	**-2.888**	-6.56
2WD Access	0.874	0.19	-	-	**0.536**	2.41	0.893	1.47	**3.633**	2.25
x Gen Hunt Exper	-0.225	-1.86	-0.149	-1.41	0.003	0.59	0.013	1.13	0.021	0.74
x HS Diploma	0.696	0.20	1.058	0.23	-0.213	-1.31	-0.337	-0.76	-0.792	-0.82
x Urban	0.105	0.05	-0.212	-0.08	**-0.233**	-2.08	**-0.671**	-2.08	**-2.104**	-2.53
x Random Effect	**5.767**	4.05	**17.64**	2.73	**0.613**	10.2	**1.981**	6.78	**4.408**	6.15
4WD Access	-	-	-	-	-0.107	-0.54	0.197	0.34	-0.683	-0.47
x Gen Hunt Exper	-	-	-	-	0.000	-0.05	0.003	0.22	0.013	0.43
x HS Diploma	-	-	-	-	**0.283**	1.99	0.527	1.23	1.475	1.59
x Urban	-	-	-	-	-0.017	-0.16	-0.067	-0.23	0.047	0.07
x Random Effect	-	-	-	-	**0.545**	9.75	**1.573**	6.92	**3.610**	5.90
No Hunters	1.591	0.30	-	-	**0.790**	3.55	**2.168**	2.98	**5.929**	2.96
x Gen Hunt Exper	0.023	0.31	0.012	0.23	**-0.014**	-3.40	**-0.041**	-3.02	**-0.105**	-2.83
x HS Diploma	-1.146	-0.31	-0.584	-0.21	0.213	1.38	0.778	1.66	1.843	1.48
x Urban	4.219	1.44	4.381	1.59	-0.009	-0.09	0.084	0.24	-0.130	-0.14
x Random Effect	**3.962**	4.42	1.662	0.56	**0.637**	12.2	**1.716**	6.71	**3.678**	5.51
On Foot	-	-	-	-	0.018	0.11	-0.425	-0.82	-0.416	-0.36
x Gen Hunt Exper	-	-	-	-	0.003	0.86	0.018	1.93	**0.056**	2.34
x HS Diploma	-	-	-	-	-0.154	-1.22	-0.719	-1.79	**-1.827**	-2.00
x Urban	-	-	-	-	0.054	0.61	0.463	1.60	0.650	1.01
x Random Effect	-	-	-	-	**0.245**	2.77	**0.952**	4.76	0.178	0.47
Forest	-1.626	-0.60	-	-	0.127	0.97	0.091	0.25	0.986	1.09
x Gen Hunt Exper	0.050	1.02	0.036	1.03	0.001	0.54	0.005	0.58	0.020	1.07
x HS Diploma	1.107	0.59	0.708	0.42	**0.233**	2.75	**0.662**	2.78	**1.385**	2.24
x Urban	-1.784	-1.34	-1.061	-0.86	-0.059	-0.71	-0.130	-0.55	-0.329	-0.70
x Random Effect	**2.825**	3.50	**8.822**	2.49	**0.384**	8.64	**1.129**	5.87	**2.227**	5.28
< 1 Moose	-4.128	-1.95	-	-	**-0.623**	-3.11	**-2.301**	-4.63	**-6.857**	-5.16
x Gen Hunt Exper	-0.085	-1.73	-0.049	-0.84	0.005	1.24	0.008	0.86	0.011	0.65
x HS Diploma	0.356	0.26	1.668	0.53	-0.229	-1.50	-0.102	-0.34	-0.481	-0.96
x Urban	**2.053**	1.98	1.730	1.39	**-0.230**	-2.30	-0.107	-0.44	0.166	0.39
x Random Effect	**4.707**	3.51	7.300	1.70	**0.577**	9.82	**1.523**	6.35	**2.780**	4.98
1 Moose	-	-	-	-	-0.069	-0.40	0.278	0.57	1.101	1.11
x Gen Hunt Exper	-	-	-	-	0.006	1.92	0.017	1.64	0.038	1.95
x HS Diploma	-	-	-	-	-0.006	-0.05	-0.055	-0.16	-0.325	-0.59
x Urban	-	-	-	-	-0.019	-0.20	-0.360	-1.25	**-1.434**	-2.20
x Random Effect	-	-	-	-	**0.200**	2.09	**1.153**	6.50	**2.607**	6.17

Boldface indicates statistical significance at the 5 percent level. All random coefficient estimates generated with 500 quasi-random draws.

Table 4 – Saskatchewan Parameter Estimates (cont.)

	RP Data Only, No ASCs		*RP Data Only, ASCs*		*SP Data Only*		*RP & SP Data, No ASCs*		*RP & SP Data, ASCs*	
	Est.	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***	***Est.***	***T-Stat***
Common Species	-5.614	-1.94	-	-	-0.169	-1.01	-0.634	-1.27	-1.702	-1.30
x Gen Hunt Exper	0.077	1.18	0.064	1.03	0.006	1.89	0.005	0.54	-0.001	-0.04
x HS Diploma	0.593	0.30	0.433	0.12	-0.192	-1.52	-0.192	-0.52	-0.383	-0.46
x Urban	**2.873**	2.07	1.605	0.86	0.014	0.15	0.462	1.68	0.777	1.16
x Random Effect	**5.370**	2.62	1.741	1.42	**0.183**	1.97	**0.703**	2.18	1.394	1.88
Unseen Species	-	-	-	-	-0.018	-0.11	-0.118	-0.23	-0.598	-0.45
x Gen Hunt Exper	-	-	-	-	-0.003	-0.98	-0.011	-1.11	-0.019	-0.80
x HS Diploma	-	-	-	-	0.170	1.40	0.564	1.44	1.548	1.47
x Urban	-	-	-	-	0.013	0.16	0.100	0.38	0.225	0.32
x Random Effect	-	-	-	-	**0.186**	2.20	**0.756**	3.18	0.700	0.40
SP Outside Dummy	-	-	-	-	**-2.377**	-3.37	**-6.574**	-3.64	**-17.95**	-3.33
x Gen Hunt Exper	-	-	-	-	0.020	1.48	0.053	0.97	0.117	0.84
x HS Diploma	-	-	-	-	-0.688	-1.25	**-2.353**	-2.16	**-6.234**	-1.98
x Urban	-	-	-	-	**-0.637**	-2.22	**-2.269**	-2.64	-4.478	-1.55
x Random Effect	-	-	-	-	**2.300**	18.21	**7.130**	7.05	**18.59**	5.65
WMZ #55 ASC**	-	-	**-19.92**	-3.18	-	-	-	-	-2.840	-1.87
WMZ #59 ASC	-	-	**-18.32**	-2.80	-	-	-	-	**-14.50**	-5.16
WMZ #60 ASC	-	-	**-8.429**	-2.40	-	-	-	-	**-3.786**	-2.90
WMZ #62 ASC	-	-	-10.12	-1.79	-	-	-	-	**2.705**	2.55
WMZ #63 ASC	-	-	**-17.25**	-2.66	-	-	-	-	**-2.900**	-3.33
WMZ #64 ASC	-	-	**-3.488**	-2.29	-	-	-	-	**-4.106**	-2.66
WMZ #65 ASC	-	-	**-5.052**	-4.66	-	-	-	-	**-3.701**	-2.32
WMZ #66 ASC	-	-	**-2.190**	-3.64	-	-	-	-	**-3.379**	-2.49
WMZ #67 ASC	-	-	**-4.628**	-3.05	-	-	-	-	-2.152	-1.71
WMZ #68 ASC	-	-	**-4.949**	-3.96	-	-	-	-	-2.332	-1.48
Heterogeneity Test P-Values										
H_0: Interact. = 0	< 0.0001		0.2779		< 0.0001		< 0.0001		< 0.0001	
H_0: Random Eff. = 0	< 0.0001		< 0.0001		< 0.0001		< 0.0001		< 0.0001	
H_0: ASCs = 0	< 0.0001		< 0.0001		< 0.0001		< 0.0001		< 0.0001	
RP/SP Pooling Test P-Values										
Models w/	***Interact., Rand. Eff., ASCs***		***Interact., Rand. Eff.***		***Some Interact., Rand. Eff., ASCs***		***Interact., ASCs***		***Rand. Eff., ASCs***	
H_0: Common RP & SP Param. Equal	< 0.0001		< 0.0001		< 0.0001		0.0022		< 0.0001	

Boldface indicates statistical significance at the 5 percent level. All random coefficient estimates generated with 500 quasi-random draws. Alternative specific constant (ASC) for WMU #507 is excluded. For “Some Interact., Rand. Eff., ASCs” specification in RP/SP pooling test section, all interactions that were not statistically significant at $\alpha = 0.05$ in the most general RP/SP model in columns 9-10 were set to zero.

Table 5 – Saskatchewan Welfare Estimates (1994 Canadian Dollars)

Specification	***Loss of Site WMZ # 59***		***Reduction in Moose Population at WMZ #59 (>3 moose/2 days to 1 moose/day)***		***Increase in Moose Population at WMZ #66 (1 moose/2 days to 1 moose/day)***	
	Mean	***Std. Err.***	***Mean***	***Std. Err.***	***Mean***	***Std. Err.***
RP Data Only – no ASCs	-$72.92	45.1	-	-	-	-
RP Data Only – ASCs	-$135.77	44.1	-	-	-	-
SP Data Only	-$163.36	15.5	-$76.83	8.44	$34.83	3.99
RP/SP Data – no ASCs	-$80.27	6.67	-$40.23	4.57	$26.57	3.68
RP/SP Data – ASCs	-$50.51	5.60	-$42.72	4.58	$80.27	14.7
Additional Results:						
RP Data – ASCs with SP Data Fill-Ins	-$110.03	35.1	-$23.83	51.2	$28.48	39.0
SP Data – ASCs from RP Data	-$117.03	10.5	-$58.52	6.17	$22.08	3.78

Welfare measures are conditional on observed choice (von Haefen, 2003); based on 3,500 simulations with the first 500 discarded as burn-in.

Appendix Table 1
Second-Stage OLS Regression Results for RP Only Models with ASCs

Parameter	***Estimate***	***Bootstrapped T-Stats***
Alberta Data (N = 13)		
Constant	**-1.252**	-2.04
Old Trail	**2.075**	3.05
No Hunters	-4.159	-4.02
Logging	0.011	0.02
< 1 Moose	-1.203	-1.44
1/2 Moose	-0.506	-0.82
3/4 Moose	-1.341	-1.58
Saskatchewan Data (N = 10)		
Constant	-4.910	-1.19
2WD Access	-1.220	-0.22
No Hunters	2.272	0.51
Forest	-1.863	-0.81
< 1 Moose	**-5.818**	-2.05
Common Species	-3.837	-1.57

Boldface indicates statistical significance at the 5 percent level.

Notes

[1] A notable exception to this statement arises when there are observed site attributes that do not vary across individuals and trips that are endogenously determined by social interactions such as congestion. In this case, second-stage estimation will be required to construct general equilibrium welfare measures that reflect endogenous resorting arising from exogenous policy shocks (see Timmins and Murdock, forthcoming).

[2] For example, in Timmins and Murdock's (forthcoming) recreation application, there are 569 sites and 16 site attributes included in the second-stage regression. Using the same data, Murdock (2006) includes slightly more attributes in the second-stage regression.

[3] In transportation choice experiment studies where alternatives correspond to transportation modes, controls for modal type (e.g., car, bus, train) are more common. See Bhat and Castelar (2002).

[4] Although the models we report assume that the RP/SP scale ratio is constant, we also estimated models where the scale ratio varied across SP choice occasions to allow for learning and fatigue effects. Although statistical tests suggested that this generalization improved model fits significantly, welfare estimates did not change qualitatively.

[5] In both SP data sets, no individual always choose the 'opt-out' alternative. Thus, the single and double-hurdle models developed in von Haefen, Massey, and Adamowicz (2005) cannot be used to address serial nonparticipation.

[6] Parameter estimates for specifications where all interaction terms were restricted to zero and where all random effects were restricted to zero are available from the authors upon request. For the tests of the joint significance of the random effects, we constructed p-values using the approach developed by Self and Liang (1987) and Andrews (2001).

[7] These results are available in the appendix. Because credible instruments were not available for this regression, we restrictively assumed that the unobserved site attributes were orthogonal to the observed site attributes.

[8] Another notable difference between our and Adamowicz et al.'s specification is the weighting scheme used for the RP and SP choices. In our specification, every RP and SP *choice* receives equal weight. In their specification, every *individual* receives equal weight. Operationally, their strategy implies weighting each RP and SP choice by the inverse of the total number of RP and SP choices, respectively. We experimented with their weighting scheme and did not find qualitatively different results from what we report in tables 2 through 5.

[9] Another explanation may be that the RP choices reflect group decisions made by multiple participants in a common recreation trip, while the SP choices reflect decisions made by a single individual.

[10] We do not consider general-equilibrium welfare measures that account for the endogenity of congestion as in Timmins and Murdock (forthcoming) although our empirical specification includes a number of congestion-like variables such as *No Hunters*, *On Foot*, and *On ATV*. We do this in part for simplicity but also because we do not have a good model for how changes in observable behavior arising from exogenous policy shocks can be linked to changes in our congestion variables.

[11] One possible approach to refining these bounds would be to iteratively search for the largest number of parameters that are consistent across the RP and SP data sets (Earnhart, 2001; Brownstone, Bunch, and Train, 2000). Given our very rich specifications of observed and unobserved heterogeneity, such an iterative approach is computationally intensive. Moreover, iteratively searching for the most general pooling specification is subject to path-dependency and other pre-test estimation errors (Leamer, 1983).

A New Approach to Value Local Recreation Sites Using Visitors' Time Allocations

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A New Approach to Value Local Recreation Sites Using Visitors' Time Allocations

Abstract

This paper applies a new revealed preference approach to valuing local recreation areas where travel costs are minimal, by relying upon the cost of time spent on site. The approach relies upon the elasticity of substitution between time spent on visiting the recreation site versus a composite commodity, and the relative time intensity of recreation versus the composite commodity. This new approach is illustrated by valuing local recreation areas along a river (Snake River) adjacent to residential areas and town. The results yield an equivalent variation estimate of welfare of \$526 per year, or \$23 per day trip. The new method appears to have promise to value heretofore difficult to value urban parks, gardens, urban forests, cultural sites, and beach recreation used by locals.

JEL classification: Q26, J20

Keywords: local recreation, revealed preference, time allocation, urban parks

Recreation Valuation Through Time Allocation rather than Expenditures

For many consumer goods the monetary marginal cost is quite small, and so it is difficult to estimate a demand curve and economic welfare arising from these goods. Examples include the Internet and outdoor recreation at nearby areas in which there is no or minimal cost to the user. However, for some of these consumer goods, the time spent using or consuming these goods is quite large. For example, according Goolsbee and Klenow (2006) the monetary cost of using the Internet is estimated to be less than 1% of income, and yet people spend between 7% and 11% of their time using the Internet. As further noted by Goolsbee and Klenow (2006), many goods used for entertainment are particularly time intensive forms of consumption. In these cases the allocation of monetary resources to the activity may not fully reflect its value to consumers. Rather, as economists since Becker (1965) have noted, the allocation of scarce time also provides useful insights on the optimizing behavior of consumers. Thus, the consumer's time allocation relative to their opportunity cost of time may provide a more complete indicator of the economic contribution that local recreation areas such as public gardens, urban parks, and cultural sites makes to the individuals utility. A rough proxy of this opportunity cost of time may be provided by the individual's wage rate. While recreation economists have devoted a great deal of study to the refining the linkage between the wage rate and the opportunity cost of time, most of the focus remains on the value of travel time and its role in valuation of recreation, see for example Englin and Shonkwiler (1995). Much less research has been devoted to the role of on-site time in valuation of recreation.

One of the first to consider the role of on-site monetary costs in estimating recreation demand was Bell and Leeworthy (1990). However, since their analysis uses per day on-site costs of beach visitors, it necessarily focuses on tourists demand, and is not applicable to local residents demand because they live so close to the site and incur little or no monetary cost per day on site. Wilman (1987) was one of the first to acknowledge the role of on-site time cost in a recreation demand model.

Recently, Goolsbe and Klenow (2006) develop a method for valuing time-intensive consumer goods by combining a relatively simple utility function, information on time shares spent on the good of interest and a composite commodity along with the an opportunity cost of time. Goolsbe and Klenow demonstrate they can estimate the demand for and welfare associated with the Internet using this approach.

The purpose of this paper is to adapt and extend their approach to using households time allocation to value nearby outdoor recreation resources that have not been amenable to valuation using standard travel cost method approaches. In particular, residents of cities or towns often have access to public gardens, city parks, cultural sites, walking paths along rivers or local beaches that are literally within walking, biking or a trivially close driving distance in which there is no or little cost to access these sites. The standard travel cost method relies upon variations in travel cost to trace out a demand curve from which consumer surplus is calculated. If there is little or no travel cost to visit the site, the demand curve may not be estimatable, or would understate the large consumer surplus realized due to the close proximity to the site. It is in these situations that this new approach of valuing these nearby recreation sites via visitors' time allocation makes an important addition to our revealed preference toolkit for valuing nearby non marketed recreational resources such as urban parks, open space, bike paths and beaches.

The Time Allocation Valuation Approach

The new approach of Goolsbee and Klenow (2006) starts with a utility function with both time and monetary expenditures on the time intensive good (e.g. recreation) and the composite commodity. The utility function involves a constant returns Cobb-Douglas household production function for producing the time intensive good and the composite commodity bundles, along with a constant elasticity of substitution utility function for preferences between the time intensive good and the composite commodity. Adapting their utility function to recreation yields:

(1) $U = \theta \left(C_r^{\alpha_r} L_r^{1-\alpha_r}\right)^{1-1/\sigma} + (1-\theta) \left(C_0^{\alpha_0} L_0^{1-\alpha_0}\right)^{1-1/\sigma}$

where (θ) is the parameter scaling the relative importance of recreation to the composite commodity.

C_r quantity of trips to the recreation site of interest,

Co quantity of all other goods, or the Composite commodity.

L_r is fraction of time spent consuming recreation at the site of interest

Lo is fraction of time spent consuming the composite commodity

α_r is the share of goods in the recreation bundle

α_o is the share of goods in the composite commodity bundle

σ is the elasticity of substitution between the recreation bundle and the composite commodity bundle.

This utility function is maximized subject to a budget constraint:

(2) $Pr*Cr + Fr+Po*Co = w(1-L_r-Lo)$

where Pr is any price for recreation, Fr is any fixed costs of recreation, Po is the "price" of the composite commodity and lower case **w** is the wage rate.

Goolsbee and Klenow also restate their set up using a more simplified utility and full income budget constraint of:

(3) $\max \theta\, Y_r^{1-1/\sigma} + (1-\theta)\,(Yc)^{1-1/\sigma}$

subject to $\lambda_r\, Y_r + F_r + \lambda o^*\, Yo = W$

where Yr and Yo are the consumption bundles of recreation and the composite commodity and λ_r and λo are the weighted average of the market prices and time costs (valued at the wage rate) of consuming the bundles Yr and Yo, respectively and capital W is Full Income (the wage rate (w) times total non-sleep time or essentially work time plus discretionary time).

Goolsbe and Klenow (2006:109) then use this simplified utility function and full income budget constraint to derive the expression for a consumer's optimal choices between the two goods as well as an expenditure function for the good of interest and the composite commodity. From the expenditure function, they derive the equivalent variation for the good of interest (the amount of money a consumer would need to provide the same utility as the opportunity to consumer the local recreation good). The equivalent variation as a percentage of full income for the good of interest (here recreation) with a linear leisure demand curve becomes:

(4) $EV/(W)= Lr/(2[\sigma\ (1-(Lr/1-(Fr/W)))]$

The annual monetary EV is calculated by multiplying the EV/W (which is a percent) times Full Income.

The values of Lr and Fr can be identified in most recreation survey data sets. For example, Lr could be calculated as the annual number of trips times the on-site time divided by leisure time.

Empirically Estimating the Elasticity of Substitution (σ)

The empirical key to this new method is to estimate σ, the elasticity of substitution between the recreation bundle and the composite commodity bundle. As noted earlier, for time intensive goods, the elasticity can be estimated using variations in the opportunity cost of time and consumption of recreation. Reliance upon the wage rate as a proxy for the opportunity cost of time is consistent with the conceptual approach of Englin and Shonkwiler and their empirical results. Thus, as the wage rate (**w**) rises, we would expect less time spent on the time intensive recreation and more time spent on the less time intensive composite commodity. Using the same utility function, the consumer's optimal allocation between recreation and the composite commodity, Goolsbee and Klenow (2006) derive an expression that can be empirically estimated and allows the recovery of σ. For local outdoor recreation that has no entrance fee, and where people can walk, ride or do a very short (5 minute) drive to the recreation site, the monetary marginal cost of a trip is treated as close zero. In addition, there may be no fixed cost of having access to the site or there could be such a small fixed cost of an annual pass, that the ratio of this fixed cost to income is so small that it can be treated as approximately equal to zero. Under these conditions Goolsbee and Klenow's (2005, 2006) original equation (3) can be simplified, and then the natural log taken of both sides, yielding an equation that can easily be empirically estimated:

$$\textbf{(5)}\ \mathbf{Ln[(1- L_r)/(L_r)] = ln(A) + (\alpha_0 - \alpha_r)(\sigma-1)\ ln(w)}$$

The dependent variable is the natural log of time on the composite commodity relative to time spent on the time intensive good of interest.

The variable **A** is short hand for the ratio of the prices of recreation and time intensity of recreation to the prices and time intensity of the composite commodity. Goolsbee and Klenow (2006: 110) argue that this ratio is constant across individuals. This is due to all individuals facing the same market prices, and being price takers. Thus it can be ignored and is reflected in the constant term of the regression.

Regressing **Ln[(1- L_r)/(L_r)]** on **ln(w)** yields the coefficient (β) on **ln(w).** To calculate **σ** , one solves equation (5) for **σ,** which is:

(6) $\sigma = 1 + [(\beta_{\ln(w)}) / (\alpha_0 - \alpha_r)].$

In Goolsbee and Klenow (2006) time valuation model, α_r is calculated as:

(7a) $\alpha_r = [(P_r\, C_r)/(w*Wh)] / [((P_r\, C_r)/(w*Wh))+((Cr*Tr)/(Wh))]$

(7b) $\alpha_0 = [(P_0 *C_0)/(w*Wh)] / [((P_0 *C_0)/(w*Wh))+ ((C_0*T_0)/(Wh))]$

where **Wh** is work hours, **Cr** is recreation trips and **Tr** is on-site plus any travel time per recreation trip, and correspondingly for the composite commodity.

Essentially, each alpha represents the ratio of the share of direct expenditures to the sum of the share of direct expenditures and share of time expenditures (see Goolsbee and Klenow, (2006: 9) for more details).

Once σ is known, EV(W) can be calculated, and then the equivalent variation in monetary terms for the time intensive good can be calculated by multiplying EV(W) by Full Income (W).

Next we empirically illustrate the method by applying the approach to a nearby recreation resource used by locals.

Illustrative Empirical Example

The local recreation area is in Jackson, Wyoming, also known as Jackson Hole. Specifically, the survey asked residents of the area their annual number of visits, length of stay and any travel costs to visit the Snake River which runs through Jackson Hole. The particular stretch of river is downstream of Grand Teton National Park, where the river runs adjacent to several housing developments. In addition there are several free public parking areas where roads and highways cross the river. The levees that run along the river provide convenient recreation opportunities that are used by residents for jogging, dog walking, walking, biking, as well as fishing from shore, or even launching rafts, canoes and kayaks. Several stretches of the river provide views of the Grand Tetons and the Jackson Hole ski area. Because many people live within walking or biking distance to the river or very short (5-10 minute) drive the most of town, there may not be sufficient variation in travel costs to apply the travel cost method. Yet, the value of recreation to local residents is a topical issue as the predominant public land management agency was under pressure to sell some of the scattered publicly accessible parcels along the river.

The data used to illustrate the application of this new method comes from a random sample of Jackson Hole, Wyoming households selected by Survey Sampling Inc. A color booklet with original cover letter and postage paid return envelope was mailed to this random sample of Jackson Hole households during September 2000. After a reminder post card and a second mailing of the survey the response rate of deliverable questionnaires was 59%.

Data Analysis

The first step in applying this new method is to use the survey data to construct the variables to run the regression equation (5) that allows calculation of sigma, the substitution of visits to this recreation site for the composite commodity. The dependent variable is constructed using Lr or the share of leisure time the individual spends on visiting the site of interest. This was calculated using the on-site time per trip (plus any travel time) times the annual number of trips. This was then divided by the amount of their leisure time. The key independent variable is the opportunity cost of that time, proxied as was done by Goolsbe and Klenow as well as many recreation economists, as the wage rate (Englin and Shonkwiler, 1995). For the purposes of this adaptation of their method we follow Goolsbe and Klenow and use the full wage rate. However, the approach can be used with suitable fractions of the wage rate typically found in the recreation valuation literature such as one-third the wage rate derived from the approach of Englin and Shonkwiler.

Since our data was not originally designed expressly for this purpose, we must also approximate the wage rate based on income divided by number in the household that work and total working hours. We then adjusted the calculated wage rate down using the percentage of non-wage income from dividends, interest and rent. Future surveys might inquire specifically about the wage rate or use a series of questions to identify the respondent's opportunity cost of time and avoid our approximations. Nonetheless, our data is sufficient to illustrate the basic approach. In addition to the wage rate variable, Goolsbe and Klenow (2006) suggest demographic variables such as age and education effects may be appropriate in this equation as well.

RESULTS

Table 1 presents the results of regressing the log of the ratio of the share of time on the composite good to the share of time on the recreation good on the log of the wage rate. The wage rate coefficient is positive, indicating those with higher opportunity costs of time, spend less time at this outdoor recreation site. The wage rate coefficient is significant at only the 10% level, due perhaps to the reliance on data from which we had to construct the wage rate based on household income, rather than eliciting the wage rate or opportunity cost of time directly. In addition the low level of significance may be due to the wage rate being less than a perfect proxy for the opportunity cost of time for those workers facing a disequilbrium in the labor market. However, the results are sufficient to illustrate the adaptation of this new method to local recreation.

(Insert Table 1).

Table 1. Regression Results

Dependent Variable: ln ((1-Lr)/Lr)

Sample: 256

Variable	Coefficient	Std. Error	t-Statistic	Probabilty
Constant	5.10277	0.7414	6.882	0.00
Ln(Wage Rate)	0.19310	0.1188	1.625	0.10
Private Raft	0.14294	0.1024	1.394	0.16
Dog Walk	-0.36869	0.1824	-2.021	0.04
Private Float Fish	-0.65461	0.1833	-3.569	0.00
Hike	-0.44673	0.2066	-2.162	0.03
Gender	0.23471	0.2089	1.123	0.26
Education	-0.02482	0.0387	-0.640	0.52
Group Size	-0.02756	0.0361	-0.762	0.44

R-squared	0.090	Mean dependent var	5.11
Adjusted R-squared	0.061	S.D. dependent var	1.34
S.E. of regression	1.303	F-statistic	3.07
Log likelihood	-426.525	Prob(F-statistic)	0.00

Where:

Wage Rate is the respondent's calculated wage rate

Private Raft is a dummy variable equal to one if the visitor's recreation activity was rafting with their own boat.

Dog Walk is a dummy variable equal to one if the visitor's recreation activity was dog walking

Private Float Fish is a dummy variable equal to one if the visitor's recreation activity was float tube fishing with their own tube and no guide.

Hike is a dummy variable equal to one if the visitor's recreation activity was hiking

Gender is a dummy variable for male/female

Education is respondent's years of education

Group Size is the number of people in the respondent's recreation group

Calculation of Equivalent Variation of Recreation

In order to calculate equivalent variation as a percentage of full income (Equation (4)), we need to calculate sigma, the elasticity of substitution between recreation and the composite commodity. Sigma depends upon the regression coefficient on log of wages in Table 1 along with the difference in the shares of time devoted to the composite commodity and recreation at this site, i.e., $(\alpha_0 - \alpha_r)$. To calculate α_r we rely upon the survey data used to calculate Lr for the regression. Specifically, the time devoted to visiting this recreation site, is arrived at by multiplying the number of trips times the time per trip (on-site time and any travel time). Across the sample, which takes which an average of 23 trips, this is 69.8 hours each year, representing a time expenditure share of .03355. The direct expenditures associated with these 23 trips is $80 per year, representing an expenditure share of .00087 of money income. As can be seen by comparing the two shares, the relative importance of visiting the Snake River in terms of time expenditure share is two orders of magnitude larger, than its relative importance as indicated by its monetary expenditure share. This suggests the need to use time as the instrument to accurately value local outdoor recreation.

To calculate α_r using equation (7a), we obtain:

$$(8)\ \alpha_r = [(\$80)/(\$91{,}976)]/[((\$80)/(\$91{,}976))+((69.8)/(2080)] = .00087/(.00087+.03355) = .0253$$

Likewise for α_0 the ratio is .361, given the expenditure share on the composite commodity is .99913 (essentially, 1-.00087). Thus $(\alpha_0 - \alpha_r)$ is .336 with rounding. Given the regression coefficient on log of wages of .193, sigma is:

$$(9)\ \sigma = 1 + [(\beta_{\ln(w)})/(\alpha_0 - \alpha_r)] = 1+[(.193)/.336)] = 1.574$$

The last variable required to calculate equivalent variation as a fraction of full income, in equation (4) is Fr/W. Goolsbee and Klenow (2005: 9) calculate Fr/W, as the expenditure share on the recreation good times (Work hours/Leisure Time). In our data this yields .00031.

Given these values of the variables, equivalent variation as a fraction of full income is therefore:

$$(10)\ EV/(W) = Lr/(2[\sigma\ (1-(Lr/1-(Fr/W)))] = .012/(2[1.574*(1-(.012/(1-.00031)))]) = .003858$$

Or essentially, .39% of full income.

In our data, total income is $91,976 but wage income is on average, 53% of this, yielding a wage rate of $23.44 an hour. When multiplied by full time (work plus discretionary time), full income is $136,492. Thus, annual EV is $526 or $22.86 per day trip. This value is quite plausible as the net economic value of visiting the Snake River because visitor access ranges from zero distance to less than a couple of miles, and the recreation experience is of high quality (i.e., views of the Tetons, bald eagles along the river, trout fishing, etc.).

Conclusions

This paper illustrates a new revealed approach for valuing nearby recreation sites used predominately by local residents, whose travel costs are either zero or so minimal that standard travel cost methods cannot be used. The method relies upon observing how visitors optimally allocate their scarce time to hours of on-site consumption of recreation at the site. In particular, we use data on visitor's allocation of time between recreation and the composite commodity as a function of their wage rate to infer an elasticity of substitution between recreation and the composite commodity. This, along with the time shares allow calculation of the welfare gained from the presence of the location recreation site (i.e., equivalent variation). For nearby recreation sites where the primary cost of consumption is the visitor's time, the opportunity cost of time paired with the allocation of time may provide a more complete indicator of the economic contribution the good makes to the individuals well being.

In our empirical example for local visitors to the sections of the Snake River running through the town of Jackson Hole, Wyoming, there was a statistically significant effect of the wage rate on the time allocation to visiting the recreation site. As such we calculated an equivalent variation for a typical local resident of continued access to these local publicly accessible sites along the river of $526 a year or $23 per day trip.

Several important extensions of this new method are possible. First, would be to design survey questions to get more directly at the wage rate and opportunity cost of time, rather than relying upon income as a proxy for the opportunity cost of time. An important refinement would be to develop a more complete and flexible model to decouple the opportunity cost of time from the wage rate.

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Using Bid Design and Anchoring Effect to Measure the Boundaries of WTP

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1. Introduction

The effects of bid design and anchoring on dichotomous choice contingent valuation results using follow-up question formats are well documented in a number of studies that illustrate how efficiency gains from double-bounded formats relative to single-bounded formats are eroded by anchoring effects (see for example Cameron and Quiggen, 1994; Herriges and Shogren, 1996; Alberini, Kanninen and Carson, 1997; Alberini, Boyle and Welsh, 2003; Alberini Veronesi and Cooper, 2005). A number of explanations are suggested, most of which presume that the responses to the two bid amounts indicate different underlying distributions for willingness to pay. This is of concern because if the first bid amount creates an anchor that affects the response to the second bid amount, welfare estimates may be biased. In this paper we present a model which suggests that bid designs generated to induce anchoring effects in multiple question formats can be used to provide information about respondents' bounds of uncertainty about preferences. Further, individual characteristics which are correlated with their level of uncertainty can be used to estimate systematic variations in the boundaries of WTP. We use a measures for bid design anchoring effects that have been given little emphasis in the literature, the mean and spread (the absolute value of the difference between two bids) and ratio of mean to spread. We conclude that potential bias in welfare estimates may be mitigated by using such a model to more completely define the possible ranges of WTP.

Herriges and Shogren (1996) investigate the role of uncertainty over preferences in generating anchoring effects. They hypothesize that people who are uncertain about their preferences interpret the first bid in a double-bounded question format as providing additional information about the value of the good in question. In their model, a prior distribution indicating true underlying willingness to pay (WTP) is recovered from first bid responses, while second bid responses represent an updated, or anterior distribution, that includes the anchoring effect. An analysis that does not take this weight of the anchoring effect into account reflects only the anterior distribution of WTP. They show that as the level of anchoring increases, efficiency gains from the double-bounded over the single bounded format decrease from 40% to approaching 0%. They include mail versions in their experimental design to test whether anchoring is a factor in first bid responses with and without a follow up question, and found anchoring was not significant at the 10% level. They note that since it is not possible to guarantee question order effects in mail versions of a double-bounded question format, all bids can influence all responses, so that the anterior/posterior interpretations of response patterns do not hold.

Alberini, Veronesi and Cooper (2005) show that the relationship among bid design, anchoring, and misspecification of the underlying WTP distribution is so complex that in certain cases it may not be possible to distinguish the individual effects of any one of these independent

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of the others. Using Monte Carlo simulations, they show that bid design dummies commonly used to test for starting point bias can instead pick up model misspecifications. They conclude that the amount of information that is conveyed by standard diagnostic tests for bid design bias may be overstated, and that "unless one is prepared to make assumptions about the form of the bias, it cannot be corrected for … [and] … without additional information beyond the responses to the bids themselves, econometric approaches to identifying and correcting for response bias do not appear to be fruitful." They conclude that more work needs to be done to investigate the role of bid design and how resulting bias is measured on the impact of starting point bias, and suggest that alternative approaches include using individual characteristics, such as respondents' views on question format (Alberini et al, 2005, p. 28).

Most models assume people know their preferences with certainty and use expected utility to determine a value when there is uncertainty about future demand or supply conditions. Their own values beyond these sources of uncertainty are typically assumed to be known with certainty. Alberini et al (2003) suggest that this assumption is too strong even for market goods, arguing that in many cases, people purchase goods in the marketplace not fully knowing how well that good will fit their needs and preferences. People who have had more experience with a brand or past purchases of a similar item may exhibit less uncertainty over their preferences for a market or non-market good. Thus past experience is a quantifiable variable which would be expected to be a predictor of uncertainty, and of the probabilistic range within which WTP is likely to be measured. All else equal, people with more experience would be expected to have a tighter range in which estimated WTP falls than people with less experience.

Most models of anchoring assume the anchoring effect is constant over all individuals, and don't vary by personal characteristics (education, familiarity with good in question, etc), but there is no reason to believe this to be the case. Specifically, as Herriges and Shogren point out, people with more uncertainty over their preferences, are likely to be more susceptible to anchoring if they perceive that the bid amounts provide them with information about the value of the good. In this paper we consider the possibility that the degree of anchoring may be systematically related to the bid design in a manner that can be predicted by the experiential context in which the individual relates to the good.

A number of methods of detection and measuring bid-design effects exist, depending on the purpose of the application. Alberini et al (2005) discuss commonly-used diagnostics for detecting bid design effects, and question the reliability and value of the information conveyed by these diagnostics when bid design, anchoring effects and uncertainty about functional form of WTP occur simultaneously. They refer specifically to dummy variables on bid sets; but their concerns are likely to apply to other common diagnostic tests for bid design effects in double bounded formats including the use of the 'other' bid amount as a right hand side variable. More recently, in the context of multiple (more than two) bid formats, Roach and Boyle (2002) and Rowe et al (1996) find that bid range spreads (truncated and non-truncated ranges) affect welfare estimates from multiple-bounded formats, while the means of the bid ranges do not affect WTP.

We propose that additional information from the use of the mean and spread between two bids as diagnostic tools to detect bid design should be further investigated. If people who are uncertain about their preferences infer additional information from a bid value, then it is not unreasonable to consider that information may also be conveyed by the relation between the two values. For example, a very wide spread might exacerbate the sense of uncertainty, so that the respondent places less information value on the bids, while a very narrow spread might suggest that there indeed is relevant information in the bid values that are proposed. On the other hand, for a given spread, the location of the mean also provides information. A very wide spread for a very high mean might leave an individual in doubt of the value of the information signaled by the bid values; while a very narrow spread for a very high mean leave an uncertain respondent with the impression that the information value is higher than if the spread were wider. In this paper we use mean, spread and the ratio of mean over spread to provide information about the nature of bid design induced anchoring effects. Finally, for people who have more experience with the good, we would expect that all of the bid design parameters would have less influence on WTP, than for less experienced respondents. Figure 1 and Table 1 illustrate these combinations of mean and spread for two different levels of uncertainty, as measured by experience with the good. We aim to test these conjectures in this paper.

2. Model

We consider a model of WTP in which we assume that it is not possible to measure WTP as a point, due to measurement instrument error, as well as respondent uncertainty about preferences. We instead focus our measurement on the probability locus within which WTP may fall. The boundaries of that locus are a function of respondent characteristics, such as experience with the good, which vary by individual. This model is similar to that of an electron orbiting a nucleus – where the measurement process itself hinders the ability to precisely determine where the electron lies. Attempts to measure where the electron may be found at a given time can only be accomplished within a range of probabilities, where that range is in part determined by characteristics of the atom. If we accept that in principle we cannot be certain of where 'true' WTP lies, we can instead concentrate on measuring the boundaries of where it lies. Determining how individual characteristics shape those boundaries becomes important in reducing measurement error.

In this case, there is no reason to believe that bid design would not affect measurement of WTP – it should. One explicit goal of any empirical study would then be to examine to what extent measurement error exists to determine the boundaries for WTP, and how those boundaries might be affected by. More experienced individuals are less likely to be affected by the bids they are presented with, so their locus contracts. If this is true, then instead of estimating a WTP as a point, economists may need to design data collection to estimate the locus of values within which WTP would be expected to fall, and to determine how that locus varies with individual experience with the good. Such an approach would indicate a bid design that tests the boundaries of where such a locus might fall over a range of experience.

3. The Data

Data were generated from a bid design based on two dollar amounts randomly and independently generated from the same distribution, which was based on data from a pilot stated preference survey. The two bids were presented to respondents sequentially in two independent dichotomous choice questions, one immediately following the other. The questionnaire was administered by mail, so the respondents saw both bid amounts at the same time. The sets of paired bids included a wide range of means and spreads (the absolute value of the difference between the two bid amounts). Means and spreads were randomly assigned in the bid design, so that narrow spreads and wide spreads are combined with high means and with low means. Bid order (high to low or low to high) was also randomized. Bid design variables are mean, spread and mean/spread.

The good was a recreational site visit, with a sample generated from an on-site survey, so all individuals had experience with at least one site visit. Respondents were asked to indicate the numbers of times they had made similar recreation trips over the previous 4 years. Respondents whose number of previous trips was above the sample mean were defined as having "expertise," those below the sample mean were defied as not having expertise. Thus expertise is a dummy variable where 1 = expertise. The null hypothesis that expertise is independent of the bid design is tested with an interaction term of expertise and bid design parameters.

A random effects probit specification is used, where the correlation coefficient indicates correlation between WTP in the first and second observations for each respondent in the panel. A number of studies use random effects specifications in multiple-bounded models. In general, a double-bounded specification restricts the two bid responses to be from the same underlying distribution for WTP. A random effects specification relaxes this restriction, allowing the correlation coefficient (ρ) to pick up the degree to which they are related. The double-bounded model is a limiting case of the random effects specification where $\rho = 1$. When the assumption that $\rho = 1$ is relaxed, "the initial and follow-up responses provide a sequence of two single-bounded intervals around the two WTP values that are more or less correlated." (Alberini et al, 1997, page 318). Alberini, Kanninen and Carson (1997) compare WTP results from random effects probit versus double-bounded specifications for three data sets to demonstrate that WTP estimates from the two specifications tend to converge with increasing correlation coefficients in the random effects specification. In their study based on three different data sets, the correlation coefficients varied from 0.36 to 0.96. As would be expected, the standard errors are lower in the double-bounded specifications in all three cases. Alberini, Boyle and Welsh (2003) use a random effects probit specification in a multiple bounded panel model with an extremely low correlation coefficient of 0.06. Using a random effects model is an effective approach when one does not wish to assume that latent WTP is identical in multiple question formats, including with data from double bounded question.

4. Results

DeShazo (2002) and others find that willingness to pay is affected by bid order (low to high or high to low). Using a dummy variable that controlled for order effects (high bid to low, or low bid to high) we were unable to reject the null that bid order has no effect on the model. It is not surprising that bid order does not affect WTP since each respondent saw the bids presented in an immediate sequence of two questions. Each bid can thereby influence the response to the other. When mean, spread and mean over spread are used to detect bid design effects; these are all positive in sign and significant.

A preliminary model used the ‘other’ offer amount as a right hand side variable as a diagnostic for detecting bid design anchoring effects. The parameter estimate was significant and positive. However, when compared with an alternative model that used mean and spread to detect bid design bias, the use of the other offer amount as a variable was a better model. This is shown as Model 2 in Table 2. Model 2 shows that the probability of a “yes” response tends to increase with the mean and spread associated with the two bids.

The next question is whether bid design effects are different for people with and without expertise with the good. Model 1 suggests that expertise by itself is not a significant predictor of the probability of a ‘yes’ response, as would have been inferred by Models 2 and 3 alone, but its interaction with the ratio of mean over spread is. Expertise is positive and significant at the 5% level in Model 3 where bid design is not taken into account. This does not change appreciably when the mean and spread of the offered bid are included in Model 2. Model 1 reveals that the action of expertise is not independent of the mean and spread and needs to be included in the model – further, expertise alone as an independent effect is no longer significant.

This result is more clearly indicated in Figure 2 where median WTP per trip is plotted for the two groups for different levels of means and spreads. Spreads are given by the values listed along the right hand side, and vary in range from s = \$3 to s= \$50. Individuals without expertise are represented in the solid lines while individuals with expertise are shown by dotted lines. Over the entire sample, an individual WTP of approximately \$390 is calculated. Using regression parameters over the sample data, WTP is calculated systematically over the sample for individuals with and without expertise for varying levels of mean and spread.

We see that the two groups behave differently with respect to how the bid design affects WTP. Both groups are less likely to be moved from the average WTP as the means of the offered bids increase for the wider categories of spread. The experienced individuals are slightly less influenced by the mean of the bids. Both groups are influenced by the bids as the spread decreases over the range of means, the extent of this influence increasing with the mean, but their behavior moves in the opposite directions. As the spread decreases, those with experience tend to undervalue the good as the mean increases. This may be explained as a tendency to mistrust the entire valuation context as the pairs of bids are seen to be very high relative to their own WTP locus. Those without expertise are more likely to be influenced by the bid amounts at narrower spreads and high means, and this influence tends to be an overvaluation. The tendency of the experienced group to be influenced by the mean changes from positive to negative as the spread increases, that is, for spreads of about \$7.77, these respondents were not affected by the mean, but for spreads above and below this amount, they were affected by the mean of the bids, but in opposite directions.

5. Conclusions

The use of both mean and spread to measure anchoring from bid design effects is effective in providing information about how uncertainty over preferences affects anchoring bias. These results are significant in that the relationship between bid design and individual characteristics may more complex than previously identified. The idea of “bias” induced by bid design has rested on the notion of a true underlying WTP as a point estimate. We might instead consider

WTP as a probability that falls within a clear and measurable set of boundaries. This would imply that data collection for estimating WTP from dichotomous choice data should incorporate contextual variables that help to delineate those boundaries. Further work includes exploring how these results might vary with different assumptions about latent WTP, and use of additional individual characteristics that may be associated with boundaries of WTP.

We conclude that further work to define boundaries of WTP loci might be a fruitful area for future work. More information about the boundaries of WTP may result in the current discussions about the extent of anchoring bias to be less relevant than discussions about how to more precisely measure boundaries. This would ultimately involve the need to include more information about respondent characteristics and definition of the context in which people develop preferences over goods, and how they reduce uncertainty over these preferences.

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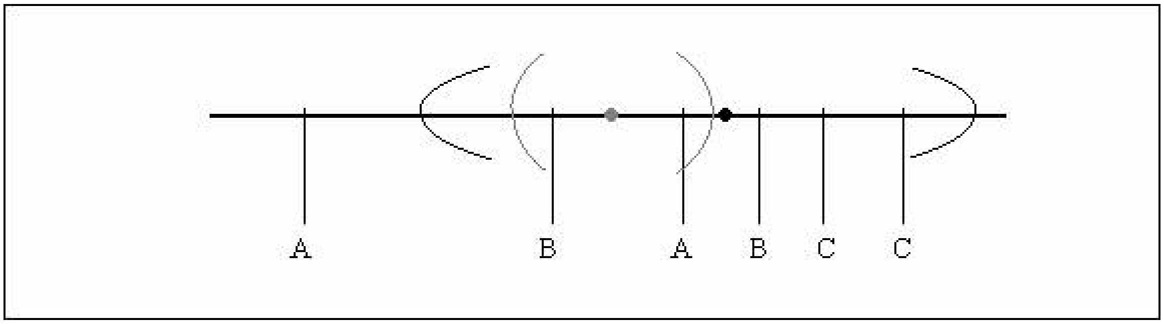
Figure 1: Behavioral model for interactions of mean and spread and experience



Wider interval (narrower parentheses) corresponds to low experience; narrower interval (rounder parentheses) corresponds to high experience. Bid pairs are indicated by letter pairs (AA, BB, CC). The relations of paired bids are described in Table 1.

Table 1: Categories of Paired Bids for Figure 1

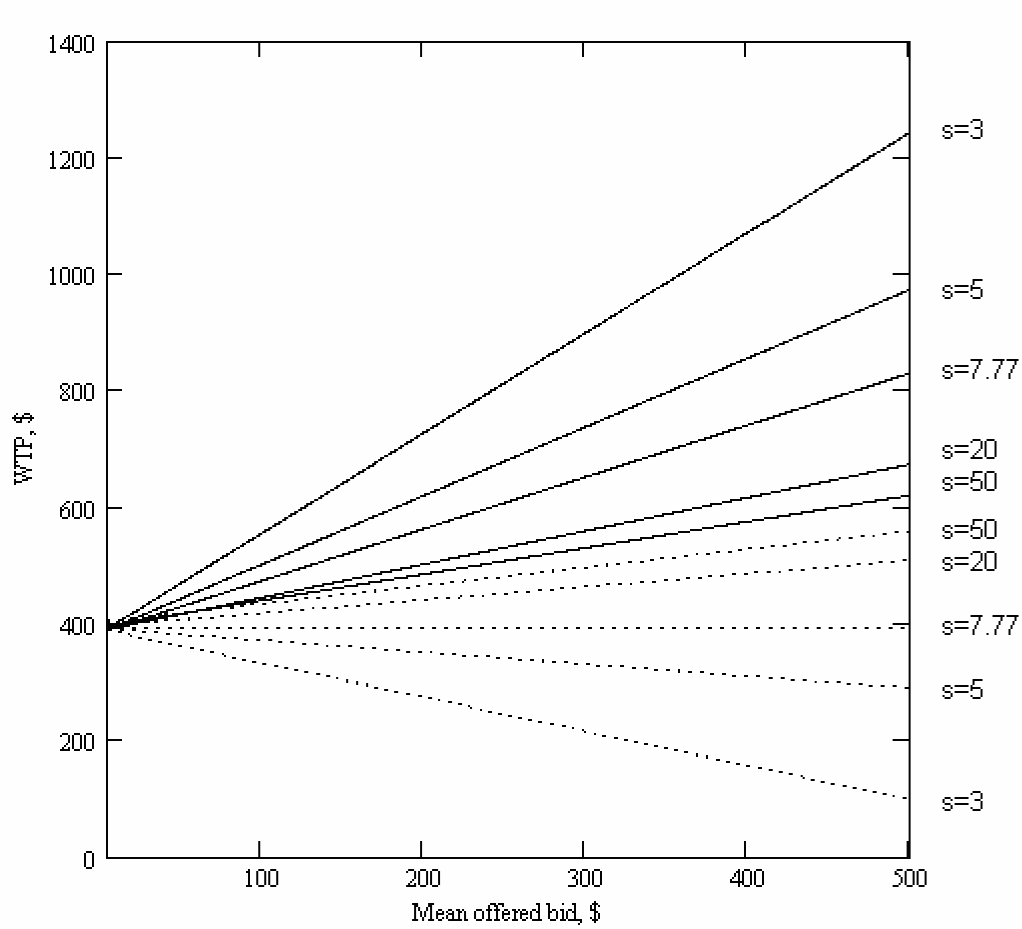
Offered Bid	Mean	Spread	Type 1	Type 2
A	low	high	N / Y	N / N
B	medium	medium	N / N	N / Y
C	high	low	N / N	Y / Y

Table 2: Random Effects Probit Model Results

	Yes/No Response	Model 1	Model 2	Model 3
	Offered bid	-0.0296 ***	-0.0295 ***	-0.0207 ***
		(0.0018)	(0.0018)	(0.0011)
	Days on site	0.5504 ***	0.5591 ***	0.5751 ***
		(0.1127)	(0.1128)	(0.0994)
	Income	0.0171 ***	0.0168 ***	0.0155 ***
		(0.0033)	(0.0033)	(0.0029)
*Site*1	Canoeing	0.1618	0.1587	-0.2569
		(0.3685)	(0.3691)	(0.3215)
	Hunt big game	2.7582 ***	2.6333 ***	2.7576 ***
		(0.6217)	(0.6221)	(0.5411)
	Rest & Relax	0.8544 **	0.8363 **	0.8748 **
		(0.4267)	(0.4280)	(0.3740)
*Site*2	Canoeing	1.8876 ***	1.8327 ***	1.5314 ***
		(0.3338)	(0.3344)	(0.2862)
	Kayaking	1.7101 ***	1.6629 ***	1.3729 ***
		(0.5999)	(0.6006)	(0.5298)
	Rest & Relax	1.2368 ***	1.1907 ***	0.9600 **
		(0.4376)	(0.4382)	(0.3823)
	Hike	1.1994 ***	1.1462 ***	0.9479 ***
		(0.3619)	(0.3623)	(0.3145)
	Car camping	1.2737 ***	1.2302 ***	0.9818 **
		(0.4395)	(0.4404)	(0.3841)
	Back-packing	1.9203 **	1.8320 **	1.5618 **
		(0.9516)	(0.9528)	(0.8388)
*Site*3	Canoeing	2.0183 ***	2.0031 ***	2.4778 ***
		(0.5461)	(0.5444)	(0.4625)
	Fishing	1.9904 **	1.8720 **	2.2807 **
		(1.0673)	(1.0434)	(0.9069)
	Rest & Relax	2.8844 ***	2.8858 ***	3.1302 ***
		(0.9423)	(0.9384)	(0.8269)
	Age	0.0167 **	0.0175 **	0.0170 ***
		(0.0070)	(0.0070)	(0.0062)
	Expertise	0.2393	-0.3058 **	-0.2615 **
		(0.3123)	(0.1738)	(0.1537)
	Mean offered bid	0.0109 ***	0.0133 ***	
		(0.0027)	(0.0022)	
	Bid spread	0.0073 **	0.0042 **	
		(0.0034)	(0.0021)	
	Mean / Spread	0.1198 **		
		(0.0575)		
	(Mean/Spread)*Expertise	-0.2047 **		
		(0.0975)		
	Constant	-1.6536 ***	-1.4228 ***	-0.4337
		(0.4790)	(0.4634)	(0.3894)
	Log likelihood	-1533.057	-1537.0101	-1569.2792
	ρ	0.8529	0.8534	0.8157
		(0.0118)	(0.0117)	(0.0127)
	Number of observations	2996	2996	2996
	Number of groups	1498	1498	1498
	Likelihood-ratio test		7.91	72.44
	(% significance level)		(98.08)	(100.00)

[a] Standard errors shown in parentheses.
* Significant at or above the 10% level.
** Significant at or above the 5% level.
*** Significant at or above the 1% level.

Figure 2: Mean and Spread impacts on WTP by level of Expertise



PART 3: Papers Supporting Objective:
"Calculate the Benefits and Costs of Agro-Environmental Policies"

A TEST OF PROXIMITY AS A PROXY FOR ENVIRONMENTAL EXPOSURE IN HEDONIC MODELS†

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ABSTRACT

Housing location affords individuals proximity to site-specific amenities (disamenities) such as parks, wetlands, airports, nuclear facilities, landfills, brownfields, and other contaminated sites. A deep vein of economic literature examines the economic effects of environmental conditions tied to location using the hedonic method. Distance to the geographic feature of interest is commonly used as a proxy for exposure to the condition. There exists little evidence regarding the degree to which proximity-based measures of net benefits reflect the value of removing the condition and thus eliminating exposure. However, it is this latter net benefit estimate that is of direct interest to policy makers. This paper reports the results of a conjoint choice experiment conducted in association with a hedonic property value study to test the role of proximity to a contaminated site as a proxy for exposure to the site, and to estimate the benefits of site cleanup.

JEL Codes: C21, Q51, R21

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A Test of Proximity as a Proxy for Environmental Exposure in Hedonic Models

Introduction

A deep vein of the environmental economics literature uses hedonic methods to analyze the effects of geographically-distributed environmental conditions on property markets. The mining of this vein began 40 years ago with the publication of Ronald G. Ridker and John A. Henning's (1967) pioneering study of air pollution's effects on residential property values. In the ensuing years, following the formalization of the hedonic method by Rosen (1974), real estate-based applications have been used to assess the market effects of a wide variety of site-specific amentities/disamenities including parks, wetlands, airports, brownfields, nuclear facilities, landfills, sports stadiums, and confined animal facilities.

This paper addresses a specific issue in hedonic property value estimation – the use of distance as a proxy for environmental exposure. With localized environmental anomalies, other locations are affected by their spatial relation to the place where the anomaly occurs. The linkage between cause and effect could be through "exposure" to the environmental condition, such as ease of access to a park or unmeasured dispersion of toxicants from a contaminated site. It could also arise from stigmatization (e.g., Dale *et al*., 1999; McCluskey and Rausser, 2003; Patunru, Braden, and Chattopadhyay, forthcoming). Whatever the linkage, proximity to the environmental intrusion is hypothesized to correlate with impact — the impact on neighboring properties exceeding that for more distant properties. The distance variable carries all of the weight of the localized environmental condition.

Policy analysis, on the other hand, is typically concerned with estimating the net benefits of improving the environmental condition of an offending site(s). To directly determine the market effects of a change in environmental quality at a site requires information on market transactions before and after the change in environmental quality. This information is rarely available since most policy analyses are prospective in nature. Thus, most studies rely on estimates of the value of increasing/decreasing proximity to a site to estimate the benefits of changing the environmental condition of the site. For proximity to accurately measure the value of environmental quality changes, a number of assumptions must be invoked that relate to distance to a site, environmental quality, and the resultant exposure for nearby properties.

In this paper, we report the results of a hedonic analysis of residential property sales that is directly linked at the household level to a choice experiment. The application examines the effect of contaminated water bodies on the residential property market. The choice experiment varies prices, the location of homes relative to the contamination, and the environmental condition. Comparisons are made between the hedonic gradients based on proximity to a contaminated site and willingness to pay estimates that disaggregate the effects of environmental quality at a site from proximity to the site. These data allow for direct tests of assumptions underlying hedonic models in this context – that the price differential due to proximity provides an accurate measure of the economic effect of changes in a location-based environmental condition.

In the following sections, we first offer a theoretical statement regarding the linkages between distance to a contaminated site, environmental quality, and the resultant exposure for nearby properties. We also present the specific propositions to be tested. Next, we describe the case study and data. We then present the estimation results and conclude with a discussion of the findings in relation to our hypotheses.

CONCEPTUAL FRAMEWORK

Hedonic Model

The hedonic context for measuring environmental cleanup values begins with the assumption that the purchase of real property reveals information about household values for site-specific amenities or disamenities. In the case of interest here, the disamenity is the potential health risks from exposure to hazardous substances that proximity of a residence to a contaminated site(s) conveys. Exposure and resulting health risks, X, are presumed to be a function of both the distance of the home to the site, D, and the environmental condition of the contaminated site, E, or $X = f(D, E)$. For ease of exposition, X is defined in such a way that positive increments in X would lead to increases in a homeowner's utility, i.e., X represents reductions in exposure. To be consistent with the definition of X, D represents distance from the contaminated site ($\partial X/\partial D > 0$), and E is an index of contamination for which increases indicate less environmental contamination and, hence, exposure ($\partial X/\partial E > 0$).

Focusing initially on consumers of housing, let the utility of the i^{th} individual be given by:

$$U(\underline{Z}, X, R, \alpha^i), \tag{1}$$

where $\underline{Z}$ is a vector of housing, neighborhood, and locational characteristics, X is environmental exposure at the home chosen, R is the numeraire good, and α^i are the relevant socioeconomic characteristics of individual i. The consumer maximizes utility subject to the budget constraint: $y = P(\underline{Z}, X) + R$, where $P(\underline{Z}, X)$ is the exogenous price schedule for housing. The consumer's optimal budget allocation is characterized by equality of the marginal rate of substitution between home attributes and the numeraire, with the price ratio:

$$\frac{U_X}{U_R} = P_X, \tag{2}$$

where $U_X = \partial U(\bullet)/\partial X$, $U_R = \partial U(\bullet)/\partial R$, and $P_X = \partial P(\underline{Z}, X)/\partial X$.

Following Rosen (1974), it is convenient to represent the consumer's problem as one of formulating an optimal bid for each house. The bid function for person i may be defined as:

$$U(y - \theta, \underline{Z}, X) \equiv U^0. \tag{3}$$

The indicator of an individual's socioeconomic characteristics, α, is omitted for brevity. Note that it is exposure at the housing site which enters utility, and not proximity to the contaminated site (D) or the environmental quality of the contaminated site (E). This assumes that there are no external effects of the contaminated site other than its conveyance of health risk to the housing unit. The plausibility of this assumption would depend on the land-use of the environmentally contaminated site. Later, we discuss how relaxing this assumption impacts our analysis.

Solving equation (3) for the bid function makes clear that bids are functions of both the utility level and income of the consumer:

$$\theta = \theta(\underline{Z}, X, U^0, y). \tag{4}$$

Substituting $\theta(\bullet)$ into (3) and optimizing with respect to changes in X implies the following condition for placing an optimal bid:

$$\theta_X = \frac{U_X}{U_R}, \tag{5}$$

where $\theta_X = \partial\,\theta(\bullet)/\partial X$ and all other partial derivatives are as defined earlier. A similar condition can be defined for the marginal bid for any housing characteristics, θ_Z. The bid function has the following properties: $\theta_X > 0$, $\theta_{XX} < 0$, $\theta_{U^0} < 0$, and $\theta_y = 1$. Furthermore, from equation (2), it is clear that a consumer will choose levels of housing characteristics such that $\theta_X = P_X$.

Figure 1 presents two bid function contours for a specific individual, where $U^1 < U^0$. The consumer faces the exogenous hedonic price function, $P(\underline{Z}, X)$, and, given his preferences and income, chooses a home at price P_0 with a level of environmental exposure of X_0. The price function in Figure 1 presumes the existence of $\overline{X}$, a level of exposure reduction beyond which there is no further incremental value in home prices. In other words, there is a level of exposure so small that further reductions do not add to housing value.

The value of an exogenous, marginal increase in X is given by θ_X or P_X. However, from a policy perspective, we are typically interested in discrete changes in environmental exposures. As described in Taylor (2002), under some conditions, the value or net benefit (NB) of a discrete change in the level of exposure from X_0 to X_1 is given by $P_1 - P_0$, and, under a variety of conditions, this price difference is an upper-bound on net benefits.[1]

The preceding discussion focuses on site-specific "exposure." This is the characteristic that is valued by consumers and conveyed to them via their choice of housing. However, from a policy perspective, we often wish to know the value of *cleanup* of a contaminated site. If we allow complete cleanup of a site to be represented by $\overline{E}$, then we wish to measure the net benefit of a change from the initial condition, E_0, to this upper-bound on quality. To link cleanup to changes in exposure, we only need $f(D, E)$ to be continuous and twice-differentiable in E and, for any D_0, there must exist an E such that $\overline{X} = f(D_0, \overline{E})$. For a consumer located at D_0 distance from the contaminated site, it is then the case that:

[1] The sum $P_1 - P_0$ represents the net benefit of a discrete change in X in a localized context when transactions costs associated with moving are assumed to be zero. If transactions costs are positive, but are not so great as to prohibit moving ($P_1 - P_0 - TC$, where TC is transactions costs) is an exact measure of net benefits. If households cannot relocate because transactions costs are prohibitively high (or there are no available houses identical to the one they owned prior to the change in the amenity, then ($P_1 - P_0 - TC$) is an upper bound on net benefits. See Taylor (2002) for a detailed discussion of how the hedonic price function may be used to measure net benefits of an amenity change.

$$U(D_0, \overline{E}, \underline{Z}, R) = U(\overline{X}, \underline{Z}, R).^{2} \tag{6}$$

And, given the following condition at the consumer's observed housing choice where $X=X_0$:

$$U(D_0, E_0, \underline{Z}, R) = U(X_0, \underline{Z}, R), \tag{7}$$

then, equations (4) and (6) also imply:

$$\theta(D_0, \overline{E}, \underline{Z}, U, y) = \theta(\overline{X}, \underline{Z}, U, y). \tag{8}$$

In this case, the value of site cleanup as evaluated by a market hedonic price function defined over exposure at each house is given by $P_{\overline{x}} - P_0$, where $P_{\overline{x}}$ is the price of housing at the upper-bound of exposure reduction. The result is illustrated in Figure 1, where $P_{\overline{x}}$ and P_0 indicate the house price evaluated at $\overline{X}$ and X_0 levels of exposure, respectively.

Empirical applications of the hedonic framework to real estate markets rarely include measures of house-level environmental exposure (X) and rarely have the information available that is necessary to evaluate the impacts of site cleanup, ΔE.[3] Evaluating changes in environmental quality requires information pre- and post-cleanup, whereas most studies are conducted only *ex-ante* as a part of prospective benefit/cost analyses.[4] As a result, most hedonic studies measure net benefits of cleanup relying on the following thought experiment. First, it is assumed there exists some distance from a contaminated site for which there is no exposure associated with that location. In other words, for any E_0 there exists a distance, $\overline{D}$, such that $\overline{X} = f(\overline{D}, E_0)$. This assumption is certainly plausible for most exposures that result from a

[2] This relationship implies the absence of non-use values in the utility function. More generally, a consumer may have non-use values for cleanup, but we would not expect these to be present in the sub-utility function associated with housing choice.
[3] The study by Gayer, Hamilton, and Viscusi (2000) represents an exception to this generalization. They differentiated general distance to a contaminated site from distance to a groundwater plume emanating from the site, calling the former a measure of stigmatization and the latter a measure of actual risk.
[4] Exceptions include Dale *et al.* (1999), McCluskey and Rausser (2003), and Zegarac and Muir (1998), who evaluated property markets before and after environmental cleanups.

contaminated site. The value of site cleanup is then assumed to be equivalent to the change in value of a property that would occur if that home was located at $\overline{D}$ instead of D_0. More formally, for any E_0, we assume there exists a $\overline{D}$ such that:

$$U(\overline{D}, E_0, \underline{Z}, R) = U(\overline{X}, \underline{Z}, R), \tag{9}$$

and given the following condition at the consumer's chosen house:

$$U(D_0, E_0, \underline{Z}, R) = U(X_0, \underline{Z}, R), \tag{10}$$

equation (9) implies:

$$\theta(\overline{D}, E_0, \underline{Z}, U, y) = \theta(\overline{X}, \underline{Z}, U, y). \tag{11}$$

It is important to note that equation (9) and (11) only hold if distance to the contaminated site has no external effects (positive or negative) other than its contribution to exposure.

If equations (8) and (11) hold, then the value of site cleanup can be evaluated by a market hedonic price function defined over distance of each house to an environmentally contaminated site. The net benefits associated with cleanup are computed by $P_{\overline{D}} - P_0$, where $P_{\overline{D}}$ is the price of housing at distance $\overline{D}$. In other words, under the conditions stated above:

$$P_{\overline{D}} - P_0 = P_{\overline{E}} - P_0 \tag{12}$$

While the preceding relationships hold theoretically under the conditions stated, data limitations or market/site conditions may violate (12). Incomplete specification of $\underline{Z}$ can result in inequality of (8) if the omitted variables are correlated with distance. For instance, unobserved locally undesirable land-uses that are spatially correlated with the site(s) of interest would mean that distance provides a biased estimate of the value of exposure. Or, spatially heterogeneous buffers between the contaminated site and housing areas could result in directional-distance being the correct specification (Cameron, 2006). Spatial lag or spatially-corrected error models will not alleviate these issues.

For most empirical applications, testing the degree to which net benefit estimates based on proximity to a specific site accurately reflect the net benefit of environmental change at the site is not possible within the context of just a hedonic property value model. This study employs data from a secondary source, a conjoint choice survey of recent homebuyers, to explore the degree to which net benefits based on distance correspond to those based on the underlying environmental condition. Parallel estimates of cleanup benefits are estimated using traditional hedonic property value model and a choice model. The context for the choice survey is presented next, followed by a discussion of the hypotheses to be tested.

Conjoint Model

An alternative approach for analyzing house purchase decisions employs a discrete choice framework, which permits direct estimation of the utility function parameters. Cropper *et al.* (1993) introduced random utility models (RUM) to real estate market analysis using conventional market data. Earnhart (2000, 2002), Braden *et al.* (2004), and Chattopadhyay *et al.* (2005) adapted the RUM model for use with survey data. However, none of these studies challenge the assumption that distance is an unbiased proxy for the external impacts of a site-based environmental condition.

Following Hanemann (1984), the utility that the i^{th} individual receives from purchasing the j^{th} home is assumed to be composed of a deterministic portion and a random error, ε, that reflects the unobservable (to the researcher) components of utility, or:

$$V_{ij} = V(y_i - P_j(\underline{Z}_j, X_j), \underline{Z}_j, X_j, \alpha_i) + \varepsilon_{ij}, \quad (13)$$

where V(•) represents the conditional indirect utility function and all other variables are as described above. Note, the above formulation focuses on choices as a function of exposure at a housing location, X. Substituting *f* (D, E) for X in the above yields:

$$V_{ij} = V(y_i - P_j(\underline{Z}_j, D_j, E), \underline{Z}_j, D_j, E, \alpha_i) + \varepsilon_{ij}. \tag{14}$$

Assuming the random component of utility is identically and independently distributed type-I extreme value, the probability that individual i chooses house j, from the set of J homes is:

$$\pi_i(j) = P(V_{ij} + \varepsilon_{ij} \geq V_{ik} + \varepsilon_{ik}; \forall\ j, k \in J)\ , \tag{15}$$

which can be translated into the conditional logit model: [5]

$$\pi_i(j) = \frac{\exp(\lambda V_{ij})}{\sum_{k \in J} \exp(\lambda V_{ik})}. \tag{16}$$

The value of a marginal change in D is given by:

$$\frac{V_D}{V_y} \text{ and } \frac{V_E}{V_y} \tag{17}$$

where $V_D = \partial V / \partial D$, $V_E = \partial V / \partial E$, and $V_y = \partial V / \partial y$. And, in the absence of an income effect, the expected compensating variation (CV) for a discrete change in D or E is calculated using the following log-sum formula (see Bockstael *et al.*, 1991):

$$E(CV) = \frac{1}{\gamma}\left[\ln \sum_{i \in J} \exp(V^1{}_i) - \ln \sum_{i \in J} \exp(V^0{}_i)\right], \tag{18}$$

where γ is the marginal utility of income, and the superscripts 0 and 1 indicate the observed utility in the base and new situations, respectively.

In the context of our choice survey, individuals are presented with a choice of two homes: their current home, and a home that is identical to their current home in all ways except for its

[5] Without loss of generality, we assume the location parameter is one (Louviere *et al*, 2000).

size, price, proximity to a contaminated site, and the condition of that site. This structure allows us to identify the household's willingness to make hypothetical trade-offs between house price, environmental conditions at the contaminated site, E, and proximity to the site, D.

Hypotheses

As described above, the choice survey has a structure that allows estimation of the value of proximity (or distance) holding environmental quality constant, and also allows estimation of the value of environmental cleanup holding proximity constant. Thus, we can employ the conjoint data to test the following hypothesis:

$$H_{01}: \quad NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\} = NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\}, \tag{19}$$

where NB^{CJ} is the net benefit (WTP) measure for ΔD or ΔE calculated with the conjoint survey results. Failure to reject the presumed equality provides evidence that the proximity discount is, at least in our application, a reasonable proxy for cleanup at contaminated sites.

While we would like to test a hypothesis similar to (19) using only hedonic data, as discussed above, that is not possible. Instead, we look to an ancillary test that examines the relationship between the net-benefit estimates arising from the two analytical methods. We explore whether, with distance held constant, the distance-based RUM estimates of welfare change approximates the hedonic estimate for elimination of the proximity discount. The hypothesis is formally stated as follows:

$$H_{02}: \quad NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\} = NB^{H}\{(\overline{D} - D_0), E_0, \underline{Z}\}, \tag{20}$$

where NB^H represents the net benefit estimate arising from a hedonic property value model based on changes in distance to a contaminated site, and all else is defined earlier. Support for H_{02} implies that the different analytical methods are estimating the same phenomenon.

For completeness, we also explore the relationship between the RUM estimates of welfare change associated with AOC cleanup and the hedonic estimate for elimination of the proximity discount. The hypothesis is formally stated as follows:

$$H_{03}: \quad NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\} = NB^{H}\{(\overline{D} - D_0), E_0, \underline{Z}\}. \qquad (21)$$

Support for (19), (20), and (21) together would be quite powerful, providing evidence of convergence validity for stated and revealed preference estimates, as well as support of the maintained assumptions in hedonic analyses that are routinely taken for granted, but rarely tested. If, on the other hand, we reject any (or all) of the above hypotheses, it may suggest a lack of convergent validity, or a failure of the maintained assumptions in the hedonic model. However, rejection may also occur due to limitations in the data and estimation methods, which we will attempt to explore as appropriate.

In summary, we first examine whether the conjoint choice experiment results are consistent with the principle that environmental exposure is fully captured in proximity discounts. Next, we compare the net benefit estimates from the choice experiment and the hedonic model to discover whether they produce consistent results. Failure to find equality between the two measures could arise from a failure of any one of a number of the underlying maintained assumptions. To the extent that we find a difference between NB^{CJ} and NB^{H}, we will explore source of the possible deviation.

APPLICATION[6]

Study Site

[6] A complete account of the case study is available in Braden *et al.* (2006)

The 1987 Amendments to the Great Lakes Water Quality Agreement between the U.S. and Canada designated 43 sites in the Laurentian Great Lakes, and their tributaries, as Areas of Concern (AOC). A common feature of these areas is the presence of toxic chemicals – notably polychlorinated biphenyls (PCBs) – known to cause cancer and neurological/developmental defects in humans and to bioaccumulate in aquatic foodwebs. The AOCs are essentially aquatic Superfund sites.[7] During the ensuing 20 years, only two Canadian sites and one U.S. site have been delisted. The remaining remedial activities on the U.S. side are expected to cost from $1.5 billion to $4.5 billion (Great Lakes Regional Collaboration, 2005). There is considerable interest in discerning whether further expenditures on cleanup will produce benefits consonant with the costs. A 35% non-federal match is required to secure federal funds, so the question of who benefits can also help identify potential sources of non-federal funding.

The Buffalo River, NY, AOC consists of a commercial harbor and a 6.2 mile segment of the river eastward from its terminus at the eastern end of Lake Erie.[8] A schematic map of the site appears in Figure 2. The eastern portion of the AOC includes a small section of a tributary. The AOC is flanked by a large industrial complex. The industrial sector is in decline and brownfield sites abound. Nevertheless, there are also private homes nearby – the 2000 Census counted 52,628 single-family homes within five miles. Our focus is on the effect of the AOC, and the potential benefits of remediation, on the market value of these private homes.

Real Estate Data

All data in our analysis relate to single family, owner-occupied home sales during the period January 1, 2002 to December 31, 2004 that occurred within five linear miles of any point

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[7] In the U.S., Superfund sites are the most toxic sites identified and eligible for remediation under provisions of the Comprehensive Environmental Response, Cleanup, and Liability Act of 1980. Many Great Lakes AOCs are closely linked to onshore Superfund sites.

[8] A more detailed description is available at: www.epa.gov/glnpo/aoc/buffalo.html.

along the Buffalo River AOC.[9] Figure 2 contains a schematic map of the study area and of the properties sold during the study period. These particular properties were chosen because they reflect recent conditions of the property markets in the target areas. The buyers presumably represent a random cross-section of housing and owner types. They would be likely to possess current knowledge about, and impressions of, the AOC. The study area encompasses most of the City of Buffalo, all of Lackawanna, and portions of Cheektowaga, Hamburg, and West Seneca. Two smaller jurisdictions, Blasdell and Sloan, also lie within the study area, but the data for the two small jurisdictions are sparse so we merge them with the data for the assessing jurisdictions, Cheektowaga and Hamburg, respectively. Table 1 provides selected census data for households in the vicinity of the Buffalo River AOC.

Several primary databases are combined to characterize homes sales in our study jurisdictions. The first data set came from local tax assessors and contains sales prices and dates and characteristics of the housing, including: a) lot size; b) square footage of improvements; c) age of primary structure; and d) miscellaneous housing characteristics. We normalized all prices to 2004 dollars using the house price index for the Buffalo-Niagara metropolitan statistical area provided by the U.S. Office of Federal Housing Enterprise Oversight. The first section of Table 2 reports names and definitions for each of the variables contained in the above categories and the first section of Table 3 reports summary statistics.

The second primary database describes spatial features of properties that sold during our study period. Of primary interest are: a) proximity of the house to the AOC; b) proximity of the house to other important location-specific amenities or disamenities such as the shoreline of Lake Erie, other rivers, railways, highway intersections, central business districts, airports, and

[9] We also collected data for multi-family residences, but we did not administer the survey to multi-family occupants so those data are not included in the current analysis.

local parks; and c) census tract, census block, and school district in which the house is located. Once again, location-related variable name and descriptions appear in tables 2 and 3.

The second section of Table 3 reports the distribution of sample properties across the jurisdictions in the study area. Forty-two percent of the properties are located in the City of Buffalo (of those, 70% are located north of the AOC). Cheektowaga and West Seneca contain another 23% and 25% of the parcels in our sample, respectively. Lackawanna and Hamburg/Blasdell contain a modest number of properties that sold during our study period; 7.5% and 2%, respectively. Overall, 47% of the sample is located north of the Buffalo River.

Using a GIS map of the Buffalo area,[10] we created variables that reflect each parcel's location relative to the features of interest. These variables are described and summarized in the last sections of tables 2 and 3. The 3.8 mile mean distance to the Buffalo CBD reflects a relative sparseness of sales in inner-city Buffalo. The mean distance between homes and streams other than the AOC is 1.6 miles while the mean distance to the AOC is approximately three miles. The latter distance does not vary much between the north and south sectors. Although not reported in Table 3, 12% of homes north of the AOC and 16% of the homes located to its south are within 1.5 miles of the AOC. The mean distance to the Lake Erie shoreline is 3.7 miles. Also of particular note is the proximity to rail corridors (mean=0.66 miles) and highway corridors (mean=0.77 miles; 0.95 miles to an interchange). These features figure prominently in the study area. The regional airport lies outside our study area, so the mean distance to properties in the sample is almost six miles.

Absent from Table 3 are 118 dummy variables included in the analysis, each one representing a census tract in which a property is located. Census tracts are designed to be

[10] Provided by John Whitney of the East Aurora, NY office, NRCS/USDA.

relatively homogeneous with respect to population characteristics, economic status, and living conditions.[11] By including census tract identifiers, our analysis indirectly captures infrastructure and demographic factors that influence home choices. These factors are thereby removed from the proximity variables, which then are free to reflect preferences for the object of the distance calculations (e.g., highways, parks, or the AOC) rather than conditions at the residence itself.

Survey Data

Based on the home sales data, we randomly chose 850 properties to receive a survey.[12] In order to achieve statistical significance at the taxing jurisdiction level, the sample was stratified to over-represent jurisdictions with fewer home sales. The survey was designed to complement the real estate market data. Three types of data were collected: 1) Information to verify current home characteristics and to assess respondent attitudes toward housing and the AOC; 2) responses to conjoint choice questions; and 3) household demographic information and reactions to the survey.

The choice questions asked respondents to imagine that additional homes had been on the market during their recent home-buying experience. Hypothetical homes were then offered to them, one by one. Respondents were asked whether, at the time of purchase, they would have preferred the hypothetical home to the home they actually bought. A representative choice question appears in Figure 3.

In order to focus respondents' attention on variables of interest and to make the choices as concrete as possible, the hypothetical homes were described as being identical to their current

[11] See http://www.census.gov/main/www/cen2000.html.

[12] The survey instruments were developed with assistance from the University of Illinois Survey Research Laboratory (SRL) and in cooperation with the Great Lakes Program, University at Buffalo.[12] Early versions were assessed by focus groups held at a public library branch in West Seneca, NY, in early 2005. Advanced versions were pretested in Spring 2005. For the final survey, respondents could either mail back a completed questionnaire or complete an equivalent instrument using the Zoomerang.com commercial survey website. Approximately nine percent of the responses were received online.

home except for selected attributes. Those attributes were chosen to focus on trade-offs between private aspects of homes and the surrounding neighborhood, particularly the AOC. In designing conjoint choice questions, the literature supports a balance between detail sufficient to make choices realistic and plausible, and simplicity that allows the choices to be readily understood and responded to (e.g., Louviere, Hensher, and Swait, 2000). Accordingly, we limited the choice questions to four attributes. All of the attributes were described in relation to the home currently owned or the neighborhood condition that existed at the time of purchase. The private home characteristic we used is the size of the residence. In virtually all hedonic studies of housing, size is highly significant in influencing price. The two other attributes were proximity of the house to the Buffalo River and the environmental condition of the river. For proximity, we asked respondents to imagine the river being closer to their home without changing other features of the neighborhood. The environmental condition was varied qualitatively, with toxic pollution increasing, decreasing, staying the same, or being eliminated.

Each of the three attributes, and the price of the hypothetical house, was allowed to take on four levels. The levels are summarized in Table 4. A choice alternative consisted of one level for each attribute. With four attributes and four levels per attribute, there are 4^4 or 256 possible combinations. Rather than examining all possible combinations, we used a fractional factorial design that varies the attributes in a manner that assures orthogonality – that is, that the design itself does not introduce correlation between variables. The orthogonal design extends to two-way interactions between the choice variables (Montgomery, 2000). Sixty-four choice alternatives resulted from the design.

For size and price, the levels are proportioned to the comparable value for the home currently owned. This has the advantage of scaling the alternatives to be realistic for each respondent while also transforming discrete "level" variables into continuous variables. For proximity, out of concern for the possibility that residents would not know the distance of their current home to the river, we used precise nominal deviations from current distance.[13] For the environmental variable, there is no obvious way to reduce it to a univariate index.

We randomly assigned the 64 alternatives into eight groups of eight alternatives. Each group was included in a version of the survey instrument. The versions differed only in the attribute combinations used in the conjoint choice questions. Table 5 summarizes the survey distribution and responses. Appendix tables A.1 to A.6 provide further insights into the survey responses and tests of their reasonableness and representativeness.

ESTIMATION RESULTS

Hedonic

The discussion and presentation of the hedonic model results focus on the relationship between sales price and proximity to the AOC. We examined several commonly-used functional forms for the hedonic price function that are consistent with our expectations regarding the general relationship between price and distance to the AOC (Braden *et al.*, 2006). Here, we report results for the following linear-log model:

$$saleprice = \alpha + \beta_1 \ln(AOC) + X\mathrm{B} + \varepsilon, \qquad \text{and } \beta>0 \tag{22}$$

[13] To deal with the possibility that a current home is already quite close to the river, the closer distances were qualified by the phrase "or next to the river if your current home is closer than…"

where *saleprice* and *AOC* are as defined in Table 2. All other housing attributes and location characteristics are subsumed in the vector X and are listed in Table 2.[14] The error term is given by ε. The parameters α, β_1, and vector B are to be estimated.[15]

The estimation results are reported in Table 6. Before discussing the results for the AOC, we briefly discuss the results for other housing attributes. The coefficient estimates for the variables describing the housing characteristics are of the expected sign and generally are statistically significant at the 5% level or better. We allow a cubic relationship between age of a home and sales price to reflect that although age typically reduces sales price, as homes become "quite old" (i.e., become "historic"), they often increase in value. The coefficient estimate for the variable indicating the number of bedrooms is negative and significant at the 10% level. This may seem counterintuitive, but we also control for square footage of the home, so this coefficient estimate reflects not only the value of adding a bedroom, but also of having smaller bedrooms (in order to keep the overall square footage the same).

Two types of location-related variables (other than *AOC*) are included in the model. The first type includes dummy variables that indicate the jurisdiction, census tract, and whether or not the property was located north of the AOC (see "Locational Dummy Variables" in Table 6). The jurisdiction not included in the model is the portion of the City of Buffalo to the north of the AOC.[16] The division of the City of Buffalo into two sectors is motivated by potentially important differences between the two areas. In particular, just north of the AOC is a substantial

[14] Other functional forms explored include those which modify equation (22) by taking the natural log of the dependent variable and/or taking the inverse of distance to the AOC (raised to a variety of powers) instead of taking the natural log. Results for these models are consistent with the model reported in (22). We report comparison results for the ln(saleprice) models or inverse models in footnotes when appropriate.

[15] We correct for an unknown form of heteroskedasticity using a Huber/White/sandwich estimator for the variance/covariance matrix.

[16] The dummy variable for Hamburg is excluded from the model because of collinearity between it and other location variables included in the model.

railway network, interstate highway, and an industrial zone. Residential density in this area is relatively sparse. The impacts of the AOC on homes in this area and further north may be different from those to the south, where there are fewer potentially confounding non-residential land uses separating residential areas and the AOC.

As indicated in Table 6, after controlling for home features and other location characteristics of the property, homes in the City of Buffalo south of the AOC, as well as homes in West Seneca and Lackawanna, sell for significantly less than homes in the City of Buffalo north of the AOC. The coefficient estimate for Cheektowaga is also negative but not significantly different from zero. The coefficient estimate for the dummy variable (*north*) indicating a home is located north of the Buffalo River is not significantly different from zero.

The second type of location-related variables included in the model captures distances between individual properties and geographic features of potential importance to homeowners (see Table 5, "Proximity Variables"). In our regression models, seven of these characteristics were interacted with a dummy variable indicating whether or not the property was north of the Buffalo River. Results are generally consistent with our expectations based on local market conditions.[17] Of particular note is the result that proximity to a hazardous waste site, other than the AOC, significantly reduces property values in both the northern and southern sectors.[18] Also, proximity to an uncontaminated stream does not significantly affect property values in either the northern or southern part of the study area, and home values actually increase with distance from the shoreline of Lake Erie, although this effect is insignificant for properties south

[17] All housing, location, and proximity variables are consistent in models in which proximity variables are entered as inverse distance instead of ln(distance). The manner in which continuous housing characteristic variables are entered do not vary across models.

[18] Significance of proximity to a HWS for properties to the south of the AOC is determined by an F-test of the null hypothesis that the sum of two coefficients are equal to zero (i.e., H_o: $\beta_{lnhws} + \beta_{lnhws*north} = 0$).

of the AOC. The majority of the shoreline is characterized by industrial or highway frontage and is not readily accessible. Under these circumstances, the negative impact of proximity to the shoreline is not surprising.

In general, in the southern sector, fewer location characteristics are significant predictors of property values. Proximity to a park, rail lines, uncontaminated streams, the airport, and the shoreline do not significantly impact property values. These findings are realistic for the geography of the area. For instance, to the south there are few rail lines in close proximity to households, which contrasts with the north. Also, the airport is located closer to the northern sector.

Turning to the variable of main interest, proximity to the AOC, we see in Table 6 that it has a statistically significant, negative impact on single-family homes located south of the Buffalo River, and a small and statistically insignificant impact on housing prices to the north of the AOC. The active rail lines, interstate highway, and industrial zones located just north of the Buffalo River appear to act as a buffer between the residential real estate market to the north and the AOC, and may overwhelm the influence of the AOC.[19]

The marginal impact of the AOC on homes to the south diminishes rapidly. For purposes of reporting, the distance variables are converted to 1/10th mile increments. The coefficient estimate for proximity to the AOC indicates that, for homes adjacent to the river (1/10th mile), values would increase approximately 8% of the mean house price per additional 1/10th mile of

[19] These results are robust to changes in the specification of the AOC variable (i.e., using the inverse of proximity to the AOC). To explore this conjecture further, we modify our definition of the "north" in our models. Recall, the area to the north of the AOC is characterized by crisscrossing rail lines, a heavily-industrial area, and there is a major highway corridor that lies between the AOC and most of our sample. However, several homes are located within a mile of the AOC to the north, but to the south of the highway and the major rail corridor in that area. We re-estimated our models assuming that the homes that lie between the AOC on the north-side and the highway/railroad corridor may be affected by the AOC like homes to the south of the AOC. Results remain unchanged. The magnitude of the coefficient estimate for the "south" is only marginally changed, and the impact on homes to the north continues to be insignificant.

distance[20]. But, just one mile from the river, the added value per additional 1/10th mile of distance is less than 1%. The gradient of marginal changes in home value is shown in Figure 4.

For comparison with the conjoint survey, we wish to estimate the realized capital loss associated with each property's proximity to the river. In terms of our earlier discussion, we estimate $(\mathrm{P}_{\overline{\mathrm{D}}} - \mathrm{P}_{\mathrm{D}_0})$. Based on equation (22), the reduction in property value for house j, located at a distance of D_0, is given by:

$$(\mathrm{P}_{\overline{\mathrm{D}}} - \mathrm{P}_{\mathrm{D}_0}) = \mathit{saleprice}\big|_{\overline{D}} - \mathit{saleprice}\big|_{D_0} = \beta_1\left[\ln(\overline{D}) - \ln(D_0)\right], \qquad (23)$$

where $\mathit{saleprice}\big|_{D_0}$ is the predicted price of house j at its actual distance from the AOC, and $\mathit{saleprice}\big|_{\overline{D}}$ is the predicted price of house j at the hypothetical "boundary" distance from the AOC. The boundary distance assumed is that for our study area: five miles from the AOC.

For a property located at the approximate mean distance from the AOC of 2.5 miles, the per-household estimated discount is \$4,692, or approximately 7% of a mean property value of \$84,619. For properties located closer to the AOC initially, say at 0.5 miles, the estimated per-household price discount is more substantial: \$15,588 or 23% of the mean property value of \$84,619. Additional computations for net benefits at different assumed initial conditions are presented in subsequent sections in the course of comparing the hedonic and conjoint results.

The calculation of the price discount only requires information on proximity of a property to the AOC. Thus, we calculate aggregate proximity discounts using all homes in the study area,

[20] A mean sale price of \$84,619 is used. This is the mean of the conjoint response sample, to which we later compare hedonic estimates. The mean sale price of homes to the south of the AOC in the hedonic sample is \$66,000.

not just those which sold during 2002-2004.[21] We calculate the price discount for each home in the study area, then sum these individual amounts. The estimated total capital losses south of the AOC, within five miles, are $118 million. These losses represent approximately 5.5% ($5,142 per home, on average) of total assessed value of the properties within five miles, south of the AOC. Note that assessed value is likely to underestimate the actual market value of the property.

Conjoint Choice

First, we estimate a relationship between the attributes included in the choice questions and the utility derived from a home. Second, based on the utility model, we derive a marginal utility of income and compute compensating variation measures of expected willingness to pay for changes in the environmental condition of the AOC. We focus initially on the effect of distance (*DIST*) to the AOC on WTP.

The estimation employs the conditional logit model. In addition to the choice variables, the specification includes an alternative specific constant (*ASC*=1 for the current home; 0 for the alternative) and socio-economic characteristics that might influence choice. The characteristics that are introduced via interactions with the attributes of housing are the number of individuals in the household (*FAMILY*) and annual per capita income (within ranges; *INCOME*). A dummy variable distinguishing respondents located north of the river from those residing to its south was included in initial models, found to be consistently not a significant predictor of choice, and is omitted. Unlike the hedonic results, the survey provided no evidence of north/south separation, a matter we revisit when presenting our results.

[21] The maintained hypothesis is that the homes which have sold during our study period are effectively a random sample of homes in the entire study area. This is reasonable if proximity to the AOC does not systematically affect which types of homes are put up for sale.

The fact that each respondent provided eight data points creates the potential for correlation. Accordingly, we used a random effects panel data estimator. After experimenting with a wide range of specifications, we found that logarithmic transformations of *HOUSE* and *DIST* best fit the data. The reasoning behind this specification is the same as for the hedonic model – incremental increases should have diminishing effects on choice.

The conditional logit specification is summarized in Table 7. The environmental conditions are distinguished using effects coding (Adamowicz *et al*., 1994, 1997),[22] and the status quo (*CURR*) is the omitted condition. Its coefficient can be recovered as the sum of the negative coefficients of the other conditions. A Wald $\chi 2$ test indicates that the model is significant at the 1% level. The log likelihood is significantly higher than the general conditional logit: (-998.38 versus -1073.82). The log likelihood test for correlation within individuals is significant at the 1% level.[23]

Because of the interaction terms, the composite effect of individual variables is difficult to discern from the estimation results. Table 8 reports the composite effects for *DIST*, *FAMILY*, and changes in environmental quality at the AOC, along with their significance levels. The composite effect of *DIST* is significantly positive when the environmental condition is *CURR*, and even more so for *ADD,* but not statistically significant for *PART* or *FULL*. These results imply that reducing or eliminating the environmental threat posed by the river would neutralize the value of distance in residential choice. The composite effect of *FAMILY* is not statistically significant for all environmental conditions.

[22] Effects coding is accomplished in this case by defining three dummy variables, D1, D2, and D3. The respective environmental conditions are represented by the following combinations of dummy variables: ADD = (1,0,0); CURR = (-1,-1,-1); PART = (0,1,0); and FULL = (0,0,1). Thus, the coefficient for *CURR* equals the sum of the negative values of the coefficients of *ADD*, *PART*, and *FULL*.

[23] Likelihood ratio test of $\rho=0$: $\chi 2$ (1)=150.88, Prob > $\chi 2$ = 0.0000. When $\rho=0$, the panel-level variance component in the error terms is unimportant and the random effect conditional logit estimator is not different from the ordinary conditional logit estimator. Here, $\rho=0$ is rejected.

The estimates in Table 7 are the basis for calculating WTP. Tables 9 and 10 report WTP estimates for changes in environmental quality at the AOC, and WTP for moving further from the AOC, respectively. In Table 9, WTP estimates are computed in two ways for comparison purposes. First, we fully enumerate the response sample to calculate mean and median utility values and then convert these values to WTP measures using equation (18). The mean and median estimate of the WTP for the sample using this approach is reported in the second and third column of Table 9. A second approach is to simply evaluate the utility function for the "mean household." In other words, the utility value is computed holding all elements of the utility function constant at the sample means (or medians) for the characteristics. WTP estimates based on this approach are reported in the right-most two columns of Table 9. Aggregate net benefits are also reported, and are calculated by multiplying the mean and median values of WTP by the number of households in the population.

As indicated in Table 9, the WTP estimates are of similar magnitudes regardless of which computational approach is used. The WTP for complete cleanup of the AOC from the current condition (*FULL*) is substantial. For the household with mean sample characteristics,[24] the WTP for full cleanup is $11,408. This estimate is somewhat more conservative than the estimates based on full sample enumeration. The WTP to avoid additional pollution at the AOC, again for the mean household, is $25,120. Note, the WTP for partial cleanup is negative, although based on a composite marginal effect that is not significantly different from zero. Thus, we do not report aggregate benefits for partial cleanup. Aggregate net benefits for full cleanup range from approximately $600 million to $700 million. Given the similarity between the estimates based

[24] The mean sample values for attributes and demographics in the model are: a household income of 67,348, a family size of 2.72 members, a current house-size of 1,483 square feet, and a distance of 2.53 miles from the AOC.

on full-sample enumeration and those based on mean-household characteristics, we focus the remaining discussion on the latter WTP estimates.

Table 9 also reports WTP estimates based on just the sample located to the south of the AOC. The per-household estimates of WTP are somewhat smaller than those based on the entire sample, and the aggregate net benefits for full cleanup are between $225 million and $260 million, less than half of the total net benefits. (Approximately 43% of the households in the study area are south of the AOC.) Estimates based on just the sample to the south of the AOC will be compared to the hedonic estimates, as only significant price impacts were found south of the AOC in the hedonic model.

Table 10 reports the net benefits for relocation, calculated for the household with mean and median characteristics. In table 10, four initial distances are considered, ranging from 0.5 miles to 3.0 miles. The net benefits are thus the additional WTP for a home located at the boundary (5 miles) as compared to the initial distance, holding environmental quality constant at the current level of contamination. For a representative household located 0.5 miles from the river, the estimate of WTP to reduce exposure through moving to the study area boundary is $5,441, or approximately 6% of the mean house price for the sample. WTP falls relatively quickly as the initial distance to the AOC is increased. For a home located initially at 2.5 miles from the AOC, the value of moving to the boundary is $1,638 or less than 2% of mean house values. Comparisons of these estimates to those for changes in environmental quality are considered next.

HYPOTHESIS TESTS

H_{01}: Equivalency of Proximity Discount to WTP for Full Cleanup in Conjoint Choices

Test: $NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\} = NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\}$

Our first conjecture is that the net benefits from moving a household to the boundary of the exposure zone is equivalent to the willingness to pay for elimination of the disamenity. This is the essential interpretation of the hedonic measure of value. We test this hypothesis using just the conjoint choice data initially. To test H_{01}, the initial environmental condition chosen, E_0, is the current condition of the AOC. An initial distance to the AOC, D_0, must also be chosen. We use four initial conditions for distance: 0.5, 1.5, 2.5 and 3.0 miles from the AOC. Table 11 reports two sets of net benefit estimates for each initial condition. The first is the WTP estimate for full cleanup of the AOC, holding distance to the AOC constant at each of the specified distances, and holding all other attributes/demographics constant at the mean value for the sample – i.e., $NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\}$. The second computes the net benefits for moving a house from its initial distance to the boundary distance (five miles), holding environmental condition constant at the current AOC contamination, and holding all other attributes/demographics constant at the sample means – i.e., $NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\}$. Table 11 also reports the t-statistic associated with the test of the null hypothesis that the two net benefit estimates are equal.

As indicated in Table 11, we cannot reject the null hypothesis of equality of net benefit for all initial distances at the 10% level. Our theoretical model – and the hedonic method for this type of application – is based on the notion that the value of obtaining a no-exposure condition for housing should be the same regardless of the means by which freedom from residential exposure is achieved – either through moving outside the impact zone (holding all else constant about the housing) or through removal of the contamination itself. However, our results thus far indicate that survey respondents revealed different values for explicit full-cleanup at a contaminated site versus moving far away from that site (and leaving its contamination unchanged).

Differences in the value of eliminating exposure may be due to upward bias in the stated value of cleanup. First, it could be that non-use values were expressed through the choices made over cleanup, even though the context was the purchase of a home which should not reflect non-use values. To test this hypothesis, we compute the estimated WTP for cleanup for a home located at the boundary of our study area, five miles. For a household with mean demographics and house attributes, the WTP for cleanup, given that their house is five miles from the AOC is $10,651 and is significantly different from zero at the 1% level. This result indicates that individuals included nonuse values in their responses or that 5 miles is not considered a “safe” distance from the AOC. The latter explanation seems implausible given the nature of the contamination.

The results may also stem from a downward bias in the stated preference for distance. We may not have fully captured the importance of distance, $\overline{\mathrm{D}}$, such that $\overline{\mathrm{X}} = f(\overline{\mathrm{D}}, \mathrm{E}_0)$. Our survey design increased proximity to the river by up to two miles, and increased distance from the AOC by up to one mile. However, our extrapolation from the mean distance, as well as distances closer to the AOC, exceed these changes and may be considered “out of sample.” Another related reason for the differences might the impact of transactions costs on responses. Although we explicitly told residents to ignore transactions costs, the current home was favored to the alternative so much so that nearly half of the hypothetical homes that were putatively “superior” in terms of price, size, and exposure were not selected. (See appendix tables A.5 and

A.6). The perception of transactions costs would increase the price differential required to persuade respondents to select a hypothetical home.

H_{02}: The Hedonic Proximity Discount Approximates the Conjoint Proximity Discount

$$\text{Test: } NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\} = NB^{H}\{(\overline{D} - D_0), E_0, \underline{Z}\}$$

Although the conjoint NB for cleanup appears to be significantly greater than the estimated conjoint net benefit for removal of exposure through increased distance, it is still of interest to compare the conjoint and the hedonic estimates of the proximity discount. Results related to this hypothesis are presented in Table 12. The net benefits of relocation in the hedonic model are approximately three times as large as the estimates from the conjoint survey. However, we cannot reject the null that the two estimates are equivalent at the 10% level. We can reject the null at 11% level, however. We return to a discussion of these results after presenting the results for the last hypothesis test.

H_{03}: The Hedonic Proximity Discount Approximates Conjoint Cleanup Net Benefits

$$\text{Test: } NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\} = NB^{H}\{(\overline{D} - D_0), E_0, \underline{Z}\}$$

Of particular interest are the results for our third hypothesis. As indicated in Table 13, the hedonic proximity discount estimates are quite similar in magnitude to the conjoint estimates for full cleanup at locations up to approximately three miles from the AOC. The similarity in the net benefit estimates for properties close to the AOC is encouraging, supporting the use of hedonic proximity NB estimates as a proxy for the value of site cleanup. The “boundary distance” at which there is perceived exposure to an environmental condition is an empirical question. Our hedonic analysis suggests that, if market transactions accurately reflect perceived exposures, the AOC has little effect on property values beyond two miles from the AOC (for properties to the south of the AOC). When comparing hedonic price discounts to stated preference values for site cleanup, we find no significant differences across the two methods within 2.5 miles of the AOC. This result suggests that hedonic models estimating net benefits through "proximity discount elimination" are capable of approximating the net benefits of site cleanup, the usual policy target.

In addition, these results provide some external validity for choice surveys, although the choice survey results appear to be less accurate at revealing the extent of the market. In other words, the marginal tradeoffs made in the surveys between cleanup, distance and house price reflect observed market transactions under particular conditions – that the households are located within the market-observed impact zone (approximately 2 miles from the AOC). However, the survey WTP estimates vary little outside the market-observed impact zone. This is akin to recent experimental evidence that suggests hypothetical choice surveys are capable of estimating marginal tradeoffs quite well, but do not accurately predict the extent of the market (Taylor, *et al*, 2006).

DISCUSSION AND CONCLUSIONS

Our theoretical model is predicated on the notion that households value reductions in toxic environmental exposures. Thus, a contaminated site that would otherwise have no external positive or negative effects on surrounding property values would be observed to generate negative impacts due its contaminated condition. If households value reductions in exposures, then the question arises of whether or not all paths that lead to a reduction in exposure are valued equally. Within the context of purchasing a home, and the site-specific amenities associated with the home, our theoretical model suggests that elimination of exposure at a home should be valued equally no matter what path leads to that elimination – whether it be cleanup of the contaminated site or moving to a "safe" distance from the site. This hypothesis critically hinges on an ability to hold all else constant about housing when eliminating exposure.

We conduct a conjoint survey that parallels a market condition in Buffalo NY to allow us to examine the aforementioned proposition. While our results are certainly not unequivocal, they do provide support for valuing site cleanup *via* hedonic estimates of proximity discounts. The converse interpretation of these results are that there is encouraging evidence that choice surveys using a property value context provide cleanup value estimates that approximate theoretically consistent values estimated from market data. However, we caution that our results also suggest that stated preferences may not accurately reveal the extent of the market – a matter of great importance when computing aggregate net benefits.

Less promising are the results related to the conjoint survey estimates of the value of increasing distance from a contaminated site. The conjoint estimates for moving to five miles from the AOC are significantly different than the conjoint estimates for the value of full cleanup, and are substantially smaller than, although not significantly different than the estimated hedonic values for distance. These results may suggest that when presented with choices between distance and cleanup in a survey context, respondents find distance a poor substitute for avoiding exposure as compared to site cleanup. Nonetheless, actual market choices related to distance from a contaminated site appear to closely parallel stated preferences for site cleanup.

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Figure 1. Hedonic Price and Bid Functions

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Figure 2. Properties Sold, and Properties Surveyed, within 5 Radial Miles of the Buffalo River Area of Concern, 2002-2004

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Figure 3. Representative Conjoint Choice Question

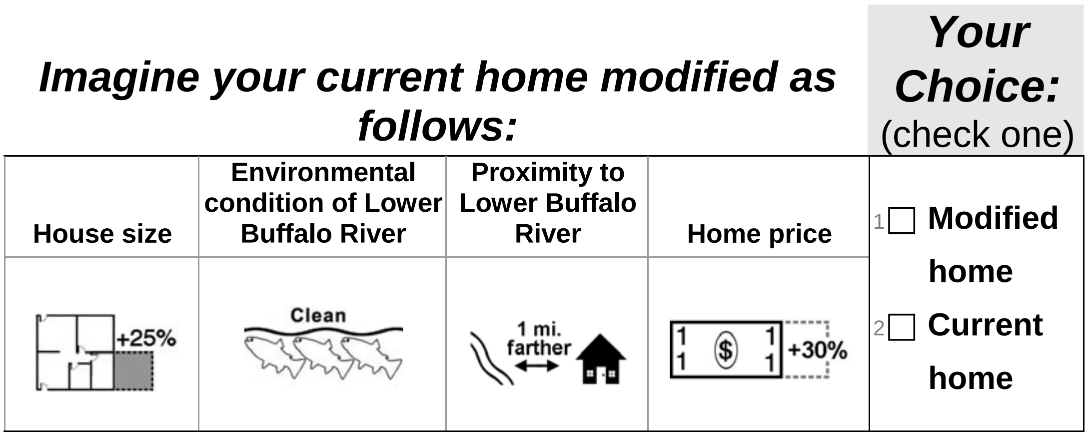

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Figure 4. Percentage Price Discounts per 1/10th mile, Owner-Occupied Dwellings South and within Five Miles of the Buffalo River AOC

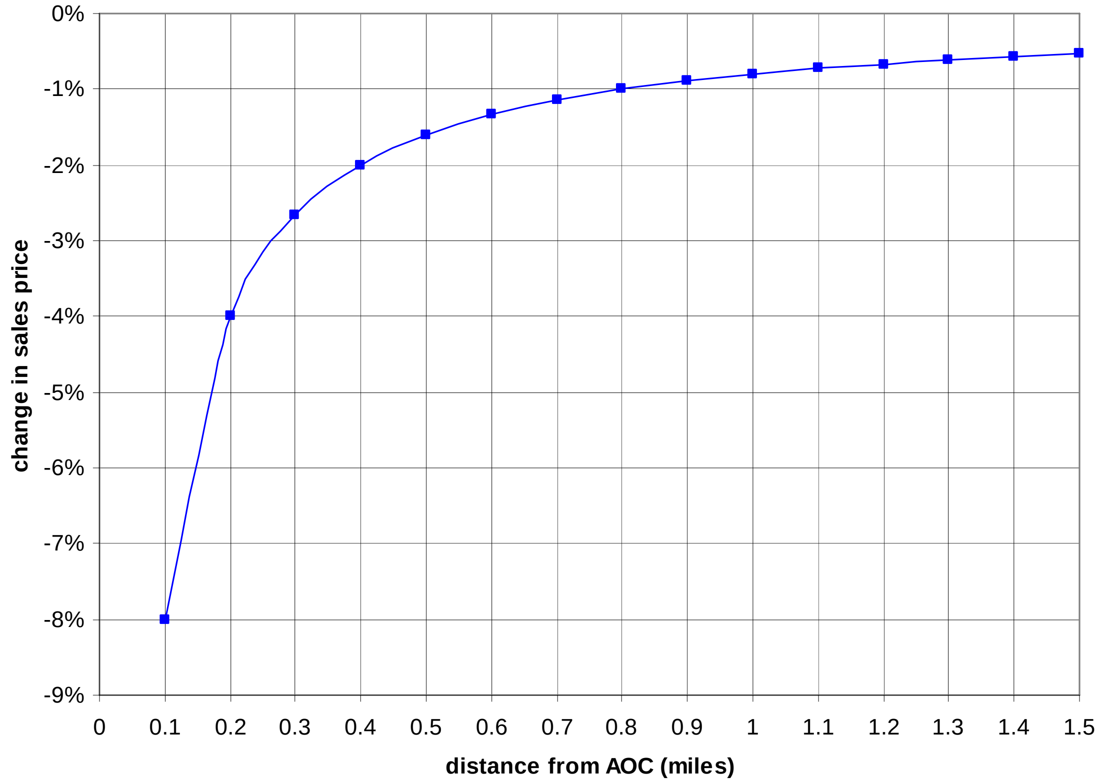

Table 1. Census Statistics for Communities Surrounding the Buffalo River AOC

	Jurisdictions near Buffalo River, NY, AOC							
Variables	Buffalo	Cheektowaga	Sloan	Hamburg	Blasdell	Lackawanna	West Seneca	Total
Population (2003)*	282,864	91,554	3,576	56,648	2,578	18,394	45,032	500,646
Median Age (years)	33.6	40.9	41.7	38.9	36.5	37.5	41.1	**36.4
Total Housing Units	145,574	41,910	1,789	22,833	1,282	8,951	18,982	241,321
Occupied Housing Units	122,720	40,045	1,680	21,999	1,201	8,192	18,328	214,165
Owner-occupied Housing Units	33,030	24,322	1,223	14,267	667	3,303	12,626	89,438
Median Value, Owner-occupied Units (2004 $)	59,300	81,800	68,600	95,700	76,600	73,600	95,200	**77,077
Average Household Size (ft^2)	2.29	2.32	2.89	2.51	2.26	2.3	2.47	**2.36
Median Household Income (1999)	30,614	38,121	29,420	47,888	43,846	29,354	46,278	**34,798

Source: U.S Bureau of the Census (2000), except as noted

* American Community Survey, U.S. Bureau of the Census (2003)

** Weighted Average of the Median

Table 2. Variable Description for Buffalo Area Hedonic Data

Housing Characteristics	
saleprice	sales price of parcel in 2004 dollars
acres (acres2)	acreage of parcel (number of acres squared)
age (age2)	age of home at time of sale (age of home squared)
sfla	square feet of living area
bedrooms	number of bedrooms
fullbaths	number of full-bathrooms
halfbaths	number of half-bathrooms
grade_ab	dummy variable =1 if tax assessor assigns a quality grade of "a" or "b" (on a scale of a, b, c, d, e, with a being the highest quality)
grade_de	dummy variable =1 if tax assessor assigns a quality grade of "d" or "e" (on a scale of a, b, c, d, e, with a being the highest quality)
grade_c	dummy variable =1 if tax assessor assigns a quality grade of "c" (on a scale of a, b, c, d, e, with a being the highest quality) – this category is omitted from the models.
cape	dummy variable =1 if home is described as a cape-cod style
colonial	dummy variable =1 if home is described as a colonial style
oldstyle	dummy variable =1 if home is described as "old-style"
otherstyle	dummy variable =1 if home is described other than the three categories listed above. Category contains mainly ranch-style homes and is omitted from the model.
fullbasement	dummy variable =1 if the home has a full basement
fireplace	dummy variable =1 if the home has at least one fireplace

Location Variables	
Buffalo_N	dummy variable =1 if the parcel is in the City of Buffalo, north of AOC
Buffalo_S	dummy variable =1 if the parcel is in the City of Buffalo, south of AOC
Cheektowaga/Sloan	dummy variable =1 if the parcel is located in Cheektowaga or Sloan
West Seneca	dummy variable =1 if the parcel is located in West Seneca
Lackawanna	dummy variable =1 if the parcel is located in Lackawanna
Hamburg/Blasdell	dummy variable =1 if the parcel is located in Hamburg or Blasdell
north	dummy variable =1 if the parcel is located north of the Buffalo River (regardless of which jurisdiction the parcel is located in)
Census Tract Identifiers	A series of dummy variables indicating the census tract in which each property is located. Parcels in our census tracts lie in 118 different census tracts – these variables are not reported in the models for succinctness.

Proximity Variables (all distances measured in miles)	
cbd	Distance to the central business district
delpark	Distance to Delaware Park , a significant park in the northern edge of the study area
park	Distance to the closest park
rail	Distance to the closest segment of a rail line
stream	Distance to the nearest stream, other than the AOC
airport	Distance to the Buffalo Airport
hws	Distance to the nearest hazardous waste site
hwy	Distance to the nearest point on a major highway
hwyx	Distance to the nearest highway interchange
shore	Distance to the shoreline of Lake Erie
AOC	Distance to the AOC

Table 3. Summary Statistics for Buffalo Area Hedonic Data

Housing Characteristics				
	Mean	Std. Dev.	Min. Value	Max. Value
saleprice	100,006.4	74,347.3	10,201	1,092,945
acres	0.184	0.241	0.015	4.583
age	60.82	29.11	0	204
sfla	1,543.43	635.08	480	9,717
bedrooms	3.22	0.81	1	10
fullbaths	1.25	0.56	1	7
halfbaths	0.34	0.49	0	5
grade_ab	0.07	0.26	0	1
grade_de	0.03	0.17	0	1
cape	0.21	0.40	0	1
colonial	0.09	0.28	0	1
oldstyle	0.41	0.49	0	1
fullbasement	0.85	0.35	0	1
fireplace	0.23	0.42	0	1
Location Variables				
	N (total=3,474)	% of total		
Buffalo_N	1,041	29.97		
Buffalo_S	427	12.29		
Cheektowaga/Sloan	794	22.86		
West Seneca	881	25.36		
Lackawanna	261	7.51		
Hamburg/Blasdell	70	2.01		
North	1,633	47.01		
Proximity Variables (Non-AOC)				
	Mean	Std. Dev.	Min. Value	Max. Value
cbd	5.01	1.71	0.27	8.22
delpark	4.93	2.71	0.02	9.41
park	0.55	0.34	0.00	2.29
rail	0.57	0.40	0.01	1.99
stream	1.57	1.71	0.01	5.68
airport	5.79	1.99	1.32	10.73
hws	0.66	0.31	0.01	1.94
hwy	0.77	0.46	0.02	2.17
hwyx	0.95	0.46	0.04	2.31
shore	3.70	1.74	0.13	7.09
Proximity to AOC				
	Mean	Std. Dev.	Min. Value	Max. Value
all properties:	3.05	1.27	0.08	4.99
north of AOC:	3.15	1.20	0.10	4.99
south of AOC:	2.96	1.32	0.08	4.86

Table 4. Home Attributes and Levels for Survey Choice Questions

	Home Size	Attribute Environmental Condition of River	Proximity to River	Home Price
Levels	+ 25%	Full cleanup	2 miles closer	+30%
	+15%	Partial cleanup	1 mile closer	+15%
	No change	No change	No change	No change
	-15%	Additional Pollution	1 mile further	-10%

Table 5. Survey Distribution and Response Summary

Jurisdiction	Jurisdiction ID	Mailed	Undeli-verable	Not Returned	Mail Response	Internet Response	Final Usable [c]
Buffalo	1	383	38	206	127	8	126
Cheektowaga[a]	2	208	8	121	71	6	67
Hamburg[b]	3	26	2	15	8	1	9
Lackawanna	4	59	4	26	28	1	26
West Seneca	5	174	11	94	60	7	66
Total[d]		850	63	462	294	23	294

[a] Cheektowaga includes Sloan.

[b] Hamburg includes Blasdell.

[c] 20 observations were lost because distance could not be estimated by GIS. In addition, three respondents did not answer any house choice questions.

[d] Total response=317. Total surveys successfully delivered= 787.
Response rate = 317/787= 40.2%

Table 6. Hedonic Price Analysis Results

Variable Name	Coefficient Estimate	Standard Error	Variable Name	Coefficient Estimate	Standard Error
Housing Characteristics			**Proximity Variables (non-AOC)**		
acres	29,357***	9,510	lncbd_n	-15,161	16,621
acres2	-8,493***	3,306	lndelpark_n	-17,458***	6,235
age	937***	294	lnpark_s	372	894
age2	-12.10***	3.96	lnrail	643	1,006
age3	0.04**	0.02	lnrail*north	2,891	2,121
sfla	28.48***	3.23	lnstream	-1,272	1,029
sfla*north	27.20***	6.11	lnstream*north	1,936	3,787
bedrooms	-2,353*	1,297	lnairport	-21,681	14,903
fullbaths	16,721***	3,056	lnairport*north	-6,822	20,810
halfbaths	6,010***	1,732	lnhws	2,660**	1,405
grade_ab	64,956***	6,739	lnhws*north	5,371	3,920
grade_de	-3,982	2,777	lnhwy	2,610	1,674
cape	-8,780***	1,647	lnhwy*north	1,908	2,846
colonial	14,255***	3,764	lnhwyx	-2,031	2,505
oldstyle	-21,287***	4,057	lnhwyx*north	-14,113***	5,333
fullbasement	2,948	1,846	lnshore	-3,658***	7,642
fireplace	4,609**	1,889	lnshore*north	26 ,668**	14,373
Location Dummy Variables[a]			**Proximity to AOC**		
Buffalo_S	-270,359*	157,679	lnaoc	6,770***	2,598
Cheektowaga	-207,798	147,340	lnaoc*north	-5,585	7,968
West Seneca	-274,997*	157,460			
Lackawanna	-269,223*	157,578		N = 3,474	
north	-170,662	266,023		R^2 = 0.8307	

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.
[a] Coefficient estimates for the 118 dummy variables indicating the census tracts in which houses were located are not reported for succinctness. The dummy variable for Hamburg is the omitted category.

Table 7. Random-Effects Conditional Logit Estimation of RUM

Variable	Coefficient	Std. Err.	z	P>z
ASC	1.1489***	0.1659857	6.92	0.00
ln*HOUSE*	2.15228	1.393722	1.54	0.12
*INCOME*x ln*HOUSE*	0.000036**	0.0000181	1.99	0.05
FAMILY x ln *INCOME*	0.519247	0.376185	1.38	0.17
ADD	-1.5735***	0.4784356	-3.29	0.00
INCOME x *ADD*	-0.000005	0.00000637	-0.86	0.39
FAMILY x *ADD*	0.165449	0.1154969	1.43	0.15
PART	0.390134	0.3520677	1.11	0.27
INCOME x *PART*	0.000001	0.00000486	0.21	0.84
FAMILY x *PART*	-0.104616	0.0921358	-1.14	0.26
FULL	1.2137***	0.3436655	3.53	0.00
INCOME x *FULL*	0.000002	0.0000046	0.54	0.59
FAMILY x *FULL*	-0.14516	0.0904646	-1.6	0.11
ln*DIST*	0.34430***	0.1068717	3.22	0.00
Ln*DIST* x *ADD*	0.4863**	0.2527134	1.92	0.05
Ln*DIST* x *PART*	-0.30866**	0.1376225	-2.24	0.03
ln*DIST* x *FULL*	-0.196546	0.1218464	-1.61	0.11
FAMILY x ln*DIST*	-0.06403**	0.0281501	-2.27	0.02
FAMILY x ln*DIST* x *ADD*	-0.08760	0.0595527	-1.47	0.14
FAMILY x ln*DIST* x *PART*	0.074000*	0.0432606	1.71	0.09
FAMILY x ln*DIST* x *FULL*	0.037559	0.0349905	1.07	0.28
Y-PRICE	0.00092***	0.0000943	9.86	0.00
INCOME x (*Y-PRICE*)	-5.42E-09***	1.14E-09	-4.75	0.00

Log likelihood = -998.38

No. obs. = 4,272
No.groups = 281
Wald $\chi^2(22) = 224.54$
Prob > $\chi^2 = 0$

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 8. Composite Marginal Effects of *DIST* and *FAMILY* by Environmental Condition, and of a Change in Environmental Condition[a]

A. DIST, holding FAMILY at the sample mean and Environmental Condition constant

Environmental Condition	Composite Coefficient	p-value	95% Confidence Boundaries Lower	Upper
FULL	0.0757	0.1082	-0.0166	0.1681
PART	0.0627	0.2429	-0.0425	0.1681
CURR	0.1238**	0.0296	0.0122	0.2354
ADD	0.1482***	0.0049	0.1268	0.7096

B. FAMILY, holding DIST at sample mean and Environmental Condition constant

Environmental Condition	Composite Coefficient	p-value	95% Confidence Boundaries Lower	Upper
FULL	3.5955	0.1867	-1.7420	8.9332
PART	3.6454	0.1786	-1.6670	8.9579
CURR	3.8093	0.1629	-1.5415	9.1603
ADD	3.8742	0.1517	-1.4227	9.1711

C. *Change in Environmental Condition from CURR, holding FAMILY and DIST constant at their sample means*

New Environ. Condition	Composite Coefficient	p-value	95% Confidence Boundaries Lower	Upper
FULL	0.5977***	0.0000	0.2353	0.9602
PART	-0.1704	0.3680	-0.5432	0.2022
ADD	-13162***	0.0000	-1.7213	-0.9112

[a] Confidence intervals computed using the Delta method (Green 2003, Oehlert 1992)
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 9. Conjoint Net Benefits for Environmental Change from *CURRENT* holding *DIST* constant, by Impact Sector[a]

	Estimates Based on Sample Enumeration		Estimates Based on Mean or Median Household	
New Environmental Condition	**Mean NB (% of Mean House Value) [Aggregate Net Benefit]**[b]	**Median NB (% of Median House Value) [Aggregate Net Benefit]**[b]	**Mean NB (% of Mean House Value) [Aggregate Net Benefit]**[b]	**Median NB (% of Median House Value) [Aggregate Net Benefit]**[b]
All Single Family Homes Within Five Miles of AOC				
ADD	-$29,874 (-32.6%) [-$1,572]	-$22,047 (-28.3%) [-$1,160]	-$25,120 (-27.4%) [-$1,322]	-$22,894 (-29.4%) [-$1,205]
PART[c]	-$3,552	-$3,321	-$3,253	-$1,951
FULL	$13,278 (14.5%) [$698]	$12,204 (15.6%) [$642]	$11,408 (12.4%) [$600]	$13,071 (16.7%) [$687]
All Single Family Homes South of the AOC and Within Five Miles				
ADD	-$24,550 (-30.1%) [-$565]	-$21,624 (-28.1%) [-$498]	-$23,369 (-28.7%) [-$538]	-$21,622 (-28.1%) [-$498]
PART[c]	-$3,326	-$3,336	-$3,201	-$3,699
FULL	$11,421 (14.0%) [$263]	$10,324 (13.4%) [$237]	$10,818 (13.2%) [$249]	$9,748 (12.6%) [$224]

[a] All values are 2004 dollars. Mean and median house values for the full response sample were $91,568 and $77,838, respectively. The comparable values south of the AOC were $81,359 and $76,894.
[b] Aggregate benefits are the per-household benefit (mean or median) multiplied by the number of households in the impact area (53,628 households in the entire study area and 23,037 south of the AOC), and are reported in millions of dollars.
[c] As shown in Table 8, Panel C, the estimated composite marginal effect for *PART* is not significantly different from zero at the 10% level, thus we do not compute percentage or aggregate benefits.

Table 10. Conjoint Net Benefits for Relocation from D_0 to $\bar{D}$ holding Environmental Condition at *CURRENT*, by Impact Sector.[a]

	Estimates Based on Mean or Median Household[b]	
Initial Distance (D_0)	**Mean NB (% of Mean House Value)**	**Median NB (% of Median House Value)**
All Single Family Homes Within Five Miles of AOC		
0.5 miles	\$5,441 (5.9%)	\$7,929 (10.1%)
1.5 miles	\$2,845 (3.1%)	\$4,146 (5.3%)
2.5 miles	\$1,638 (1.8%)	\$2,386 (3.0%)
3.0 miles	\$1,207 (1.3%)	\$1,759 (2.2%)
All Single Family Homes South of the AOC and Within Five Miles		
0.5 miles	\$5,116 (6.2%)	\$4,154 (5.4%)
1.5 miles	\$2,675 (3.2%)	\$2,172 (2.8%)
2.5 miles	\$1,540 (1.8%)	\$1,250 (1.6%)
3.0 miles	\$1,135 (1.3%)	\$921 (1.1%)

[a] All values are 2004 dollars. Mean and median house values for the full response sample were \$91,568 and \$77,838, respectively. The comparable values south of the AOC were \$84,619 and \$79,433.
[b] Compared to the overall response sample, the southern sub-sample has lower mean (\$25,569 vs. \$30,274) and median (\$23,333 vs. \$25,000) per-capita incomes, similar mean household sizes (2.76 vs. 2.72), and much larger median household sizes (3 vs. 2). These factors contribute to larger median estimates for the overall sample and larger mean estimates for the southern sub-sample.

Table 11. Comparison of Conjoint NB for Full Cleanup at Current Location to Conjoint NB for Relocation to the Impact Boundary given the Current Environmental Condition

Test of H_{01}: $NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\} = NB^{CJ}\{D_0, (\overline{E} - E_0), \underline{Z}\}$

Variable	Net Benefit (WTP) (std. dev.)[a]	t-statistic[b] (p-value)
	D_0 = 0.5 miles	
NB for *FULL*	$12,765 (4,315)	1.7023 (0.0892)
NB for Relocation to 5.0 miles	$5,441 (2,566)	
	D_0 = 1.5 miles	
NB for *FULL*	$11,756 (4,304)	2.4475 (0.0147)
NB for Relocation to 5.0 miles	$2,845 (1,341)	
	D_0 = 2.5 miles	
NB for *FULL*	$11,287 (4,464)	2.6486 (0.0083)
NB for Relocation to 5.0 miles	$1,638 (772)	
	D_0 = 3.0 miles	
NB for *FULL*	$11,120 (4,544)	2.6825 (0.0075)
NB for Relocation to 5.0 miles	$1,207 (569)	

[a] Standard errors are obtained using the Delta Method.

[b] T-statistic for the null hypothesis that the net benefit estimate for FULL equals the net benefit estimate for relocation to 5.0 miles.

Table 12. Comparison of Conjoint NB of Relocation to the Hedonic Price Discount

Test of H_{02}: $NB^{CJ}\{(\overline{D} - D_0), E_0, \underline{Z}\} = NB^{H}\{(\overline{D} - D_0), E_0, \underline{Z}\}$

Benefit Estimate Source	**Mean** [a] **(std. error)**	**t-statistic**[b] **(p-value)**
	D_0 = 0.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$15,588 (5,982)	1.6173 (0.1064)
Conjoint NB for Relocation to 5.0 miles	\$5,116 (2,476)	
	D_0 = 1.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$8,151 (3,127)	1.6178 (0.1052)
Conjoint NB for Relocation to 5.0 miles	\$2,675 (1,295)	
	D_0 = 2.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$4,692 (1,801)	1.6169 (0.1070)
Conjoint NB for Relocation to 5.0 miles	\$1,540 (745)	
	D_0 = 3.0 miles	
Hedonic NB for Relocation to 5.0 miles	\$3,458 (1,327)	1.6173 (0.1064)
Conjoint NB for Relocation to 5.0 miles	\$1,135 (549)	

[a] Net benefit estimates and t-tests are based on the smaller, southern sector for which hedonic price estimates indicated impacts of proximity to the AOC. Standard errors for conjoint estimates are obtained using the Delta Method. The conjoint estimates are evaluated for the household with mean demographic and attribute data.

[b] T-statistic for the null hypothesis that the Hedonic NB estimate for relocation to 5.0 miles is equal to the Conjoint NB estimate for relocation to 5.0 miles.

Table 13. Comparison of Conjoint NB for Cleanup to the Hedonic Price Discount

Test of H_{03}: $NB^{CJ}\{D_0, (\bar{E} - E_0), \underline{Z}\} = NB^{H}\{(\bar{D} - D_0), E_0, \underline{Z}\}$

Benefit Estimate Source	Net Benefit[a] (std. error)	t-statistic[b] (p-value)
	D_0 = 0.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$15,588 (5,982)	0.51 (0.6100)
Conjoint NB for *FULL*	\$12,107 (3,355)	
	D_0 = 1.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$8,151 (3,127)	0.67 (0.5082)
Conjoint NB for *FULL*	\$11,181 (3,289)	
	D_0 = 2.5 miles	
Hedonic NB for Relocation to 5.0 miles	\$4,692 (1,801)	1.55 (0.1212)
Conjoint NB for *FULL*	\$10,751 (3,457)	
	D_0 = 3.0 miles	
Hedonic NB for Relocation to 5.0 miles	\$3,458 (1,327)	1.89 (0.0588)
Conjoint NB for *FULL*	\$10,597 (3,544)	

[a] Net benefit estimates and t-tests are based on the smaller, southern sector for which hedonic price estimates indicated impacts of proximity to the AOC. Standard errors for conjoint estimates are obtained using the Delta Method. The conjoint estimates are evaluated for the household with mean demographic and attribute data.

[b] T-statistic for the null hypothesis that the Hedonic NB estimate for relocation to 5.0 miles is equal to the Conjoint NB estimate for FULL.

The Effects of Freer Trade on Water Use and Water Consumption

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The Effects of Freer Trade on Water Use and Water Consumption

Abstract

A two-sector, country-wide model of production, consumption, and regulation is developed to examine the linkages between water use and economic activity. The model indicates that water use is influenced by economic scale, composition, national income, openness to trade, and climatic factors. The relationships described by the model are tested using UNESCO data on water withdrawals and water consumption. Results indicate that increases in economic scale lead to increases in water use while openness to trade, capital investment, and national income reduce water use. Numerical simulations using the estimated results show that equal percentage increases in scale, personal income, and capital investment lead to increases in water use. Such increases are readily offset by increased openness to trade. At typical economic growth rates, trade liberalization has the potential to reduce average water using across countries by 3 to 5 percent in 5 to 10 years.

The Effects of Freer Trade on Water Use and Water Consumption

Studies of water use and water consumption show that economic activity is a primary determinant of global water use (Vorosmarty et al. 2000). Understanding the exact relationships between economic activity and water use, however, has proven difficult. Past trends in water use in some regions of the world appear to suggest that water use is positively correlated with income and economic scale. Water use in Asia, for example, shows a long-term positive trend with more rapid recent growth in water use occurring with more rapid economic expansion (Shiklomanov and Rodda 2003). However, water use in both Europe and North America appears to have peaked in the early 1980s despite continuing economic expansion and income growth (IWP, 2005). The mixed results for Asia, Europe, and North America cast doubts on simple trends and pairwise correlations between water use and national income. Moreover, forecasts of future water use based on such trends and correlations are poor predictors of future water use and usually overstate actual water use by substantial margins (Gleick 2000).

Inspection of detailed water use data suggests that water use is strongly influenced by the structure and composition of different economies. The most important water using sector is agriculture, accounting for two-thirds of global water use (Shiklomanov and Rodda 2003). The importance of agriculture in water use varies markedly across different regions and their different economies. In Asia, agriculture accounts for 80 percent of water use while agricultural water use in Europe is less than 40 percent of total water use. In Europe and North American, the industrial and domestic sectors account for more than 50 percent of overall water use (Shiklomanov and Rodda 2003). While the effect of economic composition stands out in the data, the effect of other social and economic factors, such as openness to trade, scale, and water regulations, need further research.

The analysis described in this manuscript examines how trade and economic structure influence water use and water consumption. The analysis draws on recent environmental economics research to develop an open-economy, country-wide model of the relationship between water use and economic activity. The model indicates that water use is influenced by economic scale, composition, the induced water regulation effects of national income, openness to trade, and climatic factors. The relationships described by the model are tested using UNESCO data on water withdrawals and water consumption. Results indicate that openness to trade, capital investment, and increases in national income are water conserving while increases in economic scale alone tend to increase water use.

The article is organized in the following manner. The next section describes the relevant environmental economics literature. The conceptual model describing water use is then developed and used to derive hypotheses. The subsequent section describes the data used in the analysis. The results section examines how country-level water use patterns vary with economic activity and tests the hypotheses derived from the conceptual model. The article ends with a brief conclusion.

Literature Review

A broad literature has developed since the early 1990s that examines the country-level relationships between economic structure and natural resource use. This literature traces its beginnings to empirical research by Grossman and Krueger (GK) (1991). GK examined the relationship between pollution levels and national income and estimated a so-called environmental Kuznets curve. While the environmental Kuznets curve turned out to lack generality (Dasgupta et al. 2002; Harbaugh, Levinson, and Wilson 2002), the economic factors identified by GK provided the foundation for improved theoretical models and substantial empirical work on the economic factors that influence the use of natural resources.

GK identified three economic factors--scale, composition, and production technique--as the central economic factors affecting natural resource use. Scale captures the idea that as an economy expands resource use increases if the expansion leaves unchanged the composition and technology of economic activity. Composition refers to the relative size of different economic sectors within an economy. Composition influences environmental resource use since different sectors use resources differently. Production technique refers to the production processes and technologies used by firms within an economy. Production technique matters since different technologies use different resources in different ways. Technologies also appear to vary widely across different countries.

GK viewed scale, composition, and technique as functions of national income. According to their analysis, national income was a direct measure of the overall scale of an economy. GK argued that composition was also related to national income, since people are likely to consume different mixes of goods at lower incomes than at higher incomes. Production techniques were viewed as responding to factor prices and environmental regulation. GK hypothesized that as income rose, citizens would seek to consume more of both market and environmental goods, putting pressure on regulators to improve environmental performance and, consequently, production techniques. The basic GK hypothesis was that the scale effect dominated as lower levels of income while at higher incomes, the income-technique effect dominated the scale effect. These countervailing scale and technique effects led to an inverted-U relationship in environmental resource use; pollution tended to rise with national income at lower levels of income, but pollution would read a peak as the scale effect was offset by the technique effect. Beyond this peak, pollution would fall with national income at higher levels of income.

GK tested the inverted-U hypothesis by examining air pollution and national income across a sample of 42 countries. Empirical results supported their hypothesis and the inverted-U relationship appeared was borne out in the data. At low levels of national income, pollution rose with income, suggesting that the scale effect dominates the technique effect at low levels of income. At higher levels of income, natural resource impacts fell with national income, implying that the technique effect dominates the scale effect at higher levels of national income. The empirical inverted-U relationship between resource damage and national income became known as an environmental Kuznets curve (Dasgupta et al. 2002). Other researchers replicatedthe

Kuznets curve relationship to varying degrees with deforestation (Ehrhardt-Martinez, Crenshaw, and Jenkins 2002; Koop and Lise 1999), a range of air pollutants (Cole and Rayner 2000), water pollution (Grossman and Krueger 1995; Lim 1997; Torras and Boyce 1998), and water consumption (Cole 2004).

Two types of problems emerged with the empirical Kuznets curve. First, the estimated inverted-U relationship between resource damage and income was fragile. Removing questionable observations and including more recent data made the inverted-U unstable, and even disappear (Harbaugh, Levinson, and Wilson 2002). Second, the emphasis on income and resource use obscured the important underlying economic relationships of scale, composition, and technique. There was a need to find specific, disaggregated ways of separately measuring the scale, composition, and income-regulation-technique effects as well as extending these relationships to include other economic factors such as trade liberalization (Dasgupta et al. 2002).

Barbier (2004) made progress with the income and regulation issues using a single sector growth model. In the single sector model, water supplied by a regulatory authority was an input into the economic processes that produce national income. The regulatory cost of supply water reduced economic output, but the increased availability of water improved the productivity of capital investments. The countervailing influences of regulatory cost and improved productivity meant that the productivity of water was high at low levels of income and low at high levels of national income. Empirical results confirmed the theoretical hypothesis and the important role of regulatory decisions on aggregate water use.

Trade liberalization is a final economic factor that appears increasingly important to incorporate in the analysis of country-level resource use (Copeland and Taylor 2004; Dasgupta et al. 2002). Openness to trade shifts production from areas with higher production cost to areas with lower production costs. Water costs are an especially important component of agricultural production so trade theory suggests that regional and global shifts in production lead to efficiency improving changes in water use. Sector-level analysis appears to confirm the water conserving effects of openness to trade. De Fraiture et al (2004) finds that trade liberalization reduced global water use in cereal production by more than 6 percent and irrigation by 11 percent, concluding that trade liberalization has a significant potential for reducing water use. Whether the de Fraiture et al results extend beyond the cereal sector awaits further research.

Copeland and Taylor (2003) and Antweiler, Copeland and Taylor [ACT] (2001) propose a model of resource use that matches the scale, composition, income-technique, and openness to trade with specific and separate variables at the country level. The analysis begins with a two sector model of production and incorporates environmental effects, consumption, regulation, and trade. The two sector model accounts explicitly for differences in economic composition and comparative advantage while also addressing scale and production technique effects. A regulatory authority sets incentives for pollution control in light of the tradeoffs between economic output and resource use. A country engages in trade comparative advantage and trade frictions such as transportation costs and trade policies. The model leads to separate empirical variables to represent each of the key economic factors; scale, composition, income-technique, and openness to trade. ACT use income only to identify the production technique effect since their model shows that higher income induces more stringent regulation of resource use.

ACT examine the empirical consequences of their model using air quality data for sulfur dioxide (SO_2) pollution. The analysis confirms that scale, composition, technique, and openness to trade may be decomposed into separate economic variables, and that these variables have distinct and economically important effects. The results show that SO_2 pollution increases with economic scale and economies with an economic composition favoring capital intensive industries. National income is associated with reductions in SO_2 pollution, consistent with a production technique effect induced by higher income and more effective regulation. Notably, openness to trade tends to reduce SO_2 in the typical country. Cole et al (2003) adapt the ACT model to three additional pollutions and find results similar to ACT except with respect to openness to trade. Cole et al find that openness to trade increases carbon dioxide emissions, reduces SO_2 and biological oxygen demand, and has no significant effect on oxides of nitrogen (Cole and Elliott 2003).

The ACT approach provides a theoretically rigorous and informative approach to examining the relationships between environmental resources and economic activity. In the next section, we adapt the ACT model to water use and to develop hypotheses about the relationship between water use and the four economic factors; scale, composition, income-technique, and openness to trade. The hypotheses are then tested using panel data for a broad spectrum of countries.

Theoretical Framework

Water provides economically valuable services both as a part of the natural environment as well as a factor of economic production. In the environment, water provides services such as habitat for fish and aquatic life, water inputs for plants, animals, and humans, transportation services, and aesthetic services. Water is a critical input for agricultural and many commercial processes. The theoretical model developed below adapts the resource use framework developed by Antweiler, Copeland and Taylor (2001) and Copeland and Taylor (2003) to an economic model of water use. The model identifies how economic and climatic factors influence water use.

A. Goods Production and Water

The analysis begins with an economy composed of a water using sector and a sector that, for simplicity, uses no water. The water using sector produces a good *x* and the non-water using sector produces the good *y*. Production technologies are constant returns to scale. The domestic price of the good produced by the water using sector is proportional to the world price,

(1) $$p = \beta p^w$$

where β represents the effect of trade frictions and p^w is world price. A country exports *x* if β is less than one and imports x if β is greater than one. Since only relative prices matter, the price of y is normalized to 1.

Water use in the water using sector is denoted by z and is proportional to the output of *x*,

(2) $$z = e(\theta, \omega)x$$

where $e(\theta, \omega)$ is the intensity of water use per unit of output, θ is the amount of water conservation done by the water using sector, and ω represents climatic factors that influence water conservation. Water conservation reduces water intensity per unit of output so $\partial e / \partial \theta < 0$. Water conservation is achieved by allocating units of output x to water conservation. The amount of output used in conservation is x_a and the amount of water conservation effort per unit of output is $\theta = x_a / x$. Some water conservation is worthwhile even when water is abundant, so $\partial e(0, \omega) / \partial \theta = -\infty$ and some water is always needed, so $\partial e(1, \omega) / \partial \theta > 0$.

A governmental regulatory authority supplies water infrastructure and charges a fee, τ, for each unit of water used in the production of output. Profits per unit of x are sales of output that is not used for water conservation minus the cost of water, $p^N = p(1-\theta) - \tau e(\theta)$. Firms choose the level of water conservation in order to maximize per unit profits. The first order conditions for maximizing profits lead to

(3) $$p = -\tau \partial e(\theta, \omega) / \partial \theta$$

so that, using equation (1), water conservation is a function of the cost of water relative to the world price of x,

(4) $$e = e(\tau / p, \omega)$$

Combining equations (2) and (4), overall demand for water use is the product water conservation per unit of product and total production of the water using sector,

(5) $$\begin{aligned} z &= e(\tau / p, \omega)x \\ &= e(\tau / \beta p^{w}, \omega)x \end{aligned}.$$

Water demand is a function of the supply price of water, the domestic price of *x*, climatic factors, and the total production of *x*. Alternatively, demand may be written as a function of world price times the trade friction parameter β in place of the domestic price.

Firms use capital, *K*, and labor, *L*, purchased at prices *r* and *w*, respectively, to produce either *x* or *y* using the appropriate constant returns to scale technology. Equilibrium on the production side of the economy satisfies equation (2), (3), and the zero profit conditions,

(6) $$\begin{aligned} &p^{N} = c^{X}(w,r), \quad 1 = c^{Y}(w,r) \\ &K = c_{r}^{X}x + c_{r}^{Y}y, \quad L = c_{w}^{Y}x + c_{w}^{Y}y \end{aligned}$$

The constraints on national income are a country's capital investment, available labor, and amount of water used in production. National income from the private sector is a function

(7) $$I = I(p^{N}, K, L)$$

The first derivatives of *I* with respect to its arguments define, respectively, the levels of output per worker supplied $I_{p^N} = x$ (Copeland and Taylor 2003).

B. Water pricing and regulation.

The supply price of z is τ and is set by the water infrastructure authority. The authority sets the price in order to maximize welfare given national income opportunities, the cost of supplying water, and the environmental costs of using water. The cost of supply water reduces national income, $I(p^{N}, K, L)$, by an amount $c(z,\omega)$ that increases with water supplies and varies with climatic factors, ω.

The regulatory authority sets the water price policy to maximize the well-being of the population. Well-being is represented by a welfare function that increases with income and declines linearly in water use due to environmental costs, $u\left[I + \tau z - c(z,\omega)\right] - z$ where

$I+\tau z-c(z,\omega)$ is income adjusted for revenues received by the regulatory authority and the costs of water supply. The first order condition maximizing welfare relative to the choice of the water price τ^* is

$$(8) \qquad u\left[G_x + z + \tau\frac{dz}{d\tau} - c'\frac{dz}{d\tau}\right] - \frac{dz}{d\tau} = 0$$

Equation (9) simplifies to

$$(9) \qquad \tau = c' + 1/u'$$

Equation (11) describes the water supply price. The supply price is determined by two terms, the marginal cost of water supply in terms of forgone national income, c', and the marginal environmental cost of water supply, $1/u'$. The marginal cost of forgone national income is increasing in the amount of water supplied and also varies with climatic factors. The environmental cost is the inverse of marginal utility, so environmental costs and water price are higher the smaller is marginal utility. Marginal utility of income is conventionally thought to be greater at lower levels of income and smaller at higher levels of income. The latter implies that the subjective evaluation of environmental costs is lower at lower levels of income and higher at higher levels of income. Thus, one might expect that lower income countries set a lower water price due to a lower evaluation of environmental costs while higher income countries set a higher water price due to a higher evaluation of environmental costs. Other things equal, the latter observation suggests that income has a negative impact on water use.

To make explicit the economic factors that determine it, water price is written as a function of the predetermined variables that influence $c' = c'(z,\omega)$ and $u^{i}{}' = m(\tau,\beta,p^w,I)$,

$$(10) \qquad \tau = \tau(\beta, p^w, I, \gamma)\,.$$

C. Determinants of Water Use

Equation (5) describes the demand for water and equation (10) describes the supply price of water. These two equations are combined to derive a reduced form equation describing the determinants of water use. Substituting the water price function into the water conservation function results in $e = e(\beta, p^w, I, \omega)$. Substituting the latter result into equation (5) leads to the equation for the equilibrium quantity of water use,

(11) $$z = e(\beta, p^w, I, \omega)x .$$

A country's production of the water using commodity *x* may also be written as a function of the overall output scale of the economy, *S*, and the share, φ, of the output attributable to *x* (Antweiler, Copeland, and Taylor 2001). By the Heckscher-Ohlin theorem, the composition of a two sector economy and share of total output, *S*, attributable to x is a function of a country's capital per worker, $k = K / L$, so $\varphi = \varphi(k)$. An alternative form of equation (11) is therefore

(12) $$\begin{aligned} z &= S\varphi e \\ &= S\varphi(k)e(\beta, I, \omega) \\ &= z(S, k, \beta, I, \omega) \end{aligned}$$

where the world price, p^w, since it remains constant across countries while openness to trade factor, β, varies. The first line of equation (12) confirms the GK hypothesis for water use; water use is influenced by economic scale, composition, and production technique. In this case, production technique is the technology of water conservation. The second lines and third lines define a reduced form equation that describes a set of empirically measurable determinants of water use.

Equation (12) suggests hypotheses about how water use may be influenced by its economic determinants. Water use is proportional to scale so water use is expected to rise and fall with economic scale. Water use is expected to be smaller in a country with a smaller economy and larger in a larger economy.

An increase in capital per worker and a shift toward manufacturing is likely to reduce water use. Agriculture is a major water using sector and is less capital intensive than economic sectors such as manufacturing. Also, capital investments in agriculture, such as drip irrigation and other water saving investments, are likely to reduce water use.

A reduction in trade frictions, β could result in either an increase or reduction is water use depending on a country's comparative advantage in agriculture or in non-water and less water intensive sectors. However, on average across a number of countries, a reduction in trade barriers is likely to shift production to areas with the respective comparative advantage, including a comparative advantage in water costs and pricing. The latter result is suggested by the recent results where freer trade in cereal production reduced global water use by 6 percent. We expect this result to hold generally within economies so that water use is likely to decline and water conservation is likely to increase with openness to trade.

Income enters the equation (12) through its effect on water pricing. An increase in income tends to raise the regulatory price of water. A higher price reduces water use by increasing the level of water conservation. As the country-level value of water rises, one would expect a greater investment in water conserving institutions and technology. Accordingly, increases in income are expected to reduce water use, holding other variables unchanged.

Equation (12) suggests no definite hypotheses for climatic factors. Climatic factors enter as determinants of water use through both water price and the technology of water conservation, with each having countervailing effects on water use. For instance, an increase in temperature may increase the costs of providing water and lead to a higher implicit regulatory price of water. At the same time, increased temperatures may increase the productivity of water in agriculture and lead to reduced conservation. The cost effect of temperature tends to reduce quantity demanded use while the productivity effect tends to increase quantity demanded. The net effect may only be measured empirically.

Notably, the theoretical concepts summarized in equation (12) place no restrictions on the overall functional form connecting water use with the predetermined factors. Equation (12) may have an inverted-U relationship with income, but there is nothing in the theory requiring equation (12) to be an inverted-U in income. The only hypothesis about income is that it reduces water use due to its effect on regulation and water price policy. Unlike the GK approach, scale and composition effects enter the analysis through their own, distinct variables, *S*, and *k*, so there is no implicit netting out of the scale and income-technique effects. Scale, composition, income-technique, and openness are each represented separately and each is tested separately for its effect on water use.

Data

Testing the hypotheses derived from the model requires data on water use, economic factors, and climatic variables. This section describes the data sources and measurement issues as well as the variables used to estimate the model.

The typical empirical measures of water use are water withdrawals and water consumption (Shiklomanov and Penkova 2003). Water withdrawals measure the amount of water diverted from water resources for residential, municipal, industrial, and agricultural uses. Most water withdrawn is returned to the terrestrial hydrological system in some form. For instance, water withdrawn from a river may be used for industrial cooling and then is discharged back into the river as warmer water. Water consumption refers to water that is removed from a country’s available water resources and not returned. Most water consumption is incurred through evaporation, but consumption also accounts for leakages from distribution systems and water that forms part of a finished manufactured or agricultural product (Shiklomanov and Penkova 2003). Given their distinct definitions, withdrawals and consumption are likely to have distinct economic and environmental consequences. Hence, both water withdrawals and water consumption are used as dependent variables in the empirical analysis.

The primary sources for water use data are the AQUASTAT database maintained by the Food and Agriculture Organization (2005), the World Resource Institute summary database EarthTrends (2005), and the International Hydrological Programme (IHP) database maintained by UNESCO (2005). AQUASTAT focuses primarily on the water use in the agricultural sector and the availability groundwater and surface water. AQUASTAT includes country-level water withdrawals, but has significant missing data problems for most countries. EarthTrends summarizes water resource, withdrawal, and infrastructure data from multiple sources, including AQUASTAT. EarthTrends data are missing for many countries and the data that is available is often listed for different years across different countries. Furthermore, using data collated from multiple sources raises its problems with inconsistencies in the data estimation methods used by the original sources.

The IHP database contains consistently measured data on water resources, withdrawals, and consumption. The data set focuses on regional summaries, but data are available for a large number of selected countries. The data are compiled for a consistent set of years, covering country-level annual estimates for 1900, 1940, 1950, 1960, 1970, 1980, 1990, and 1995. The data are based on a comprehensive review of original sources and alternative estimates. Withdrawal and consumption data are based on consistent definitions and protocols.

The IHP data were used in formulating the dependent variables in our analysis since the data contained both country-level withdrawals and consumption for a consistent pattern of years. To match the economic and demographic data available, use of the IHP data was limited to 1970, 1980, and 1990. Per capita withdrawals and consumption were calculated from the IHP data by dividing total water withdrawals and consumption by country-level population. Population data were obtained from the Penn World tables 5.6 (Heston et al, 2002). The dependent variables in were therefore per capita water withdrawals and consumption per capita by country for 1970, 1980, and 1990.

Country-level economic data are compiled in databases such as the World Development Indicators (TWB, 2005) and EarthTrends (WRI, 2005). However, these databases exhibit temporal inconsistencies and missing data problems, particularly in terms of matching the available economic information with water use for 1970, 1980, and 1990. In addition, most economic databases focus on national account variables such as gross national product and income, and do not contain data on capital stocks or capital per worker. The exception is the Penn World Table (PWT), version 5.6 (Heston, Summers, and Aten 2002). The PWT includes data sufficient to formulate each of the economic variables indicated as important by the theoretical analysis and the data are available for 1970, 1980, and 1990 to match the IHP water use data.

The PWT data were used to formulate measures of the four economic variables identified by the theoretical analysis, scale, capital-labor ratio, income, and openness to trade. Following ACT, gross domestic product (GDP) was used as the measure of economic scale and gross national income (GNI) was used to measure of national income. GDP and GNI are often thought of as similar, but GDP measures the total product produced within a country's borders while GNI measures the income received by a country's residents wherever it may be produced. The capital-labor ratio was measured by capital stock per worker and openness to trade was measured by total exports and imports divided by GDP (Heston 1991). The scale and income variables were both divided by population to convert them to scale and income per capita. Scale, income, and the capital-labor ratio were adjusted to U.S. dollars at the first quarter, 2005 price level.

Five climatic variables were included in the empirical analysis. Mean annual precipitation and temperature for 1970, 1980, and 1990 were derived from data compiled by the Global Historical Climatology Network (Vose et al. 1992). The latter data were available as monthly precipitation and temperature values across the reporting stations in a given country. Country level annual means were computed by averaging the monthly data for stations with a given country and year. Three climatic dummy variables were also included to denote specific climatic features contained in standard climatic descriptions (CIA, 2005). A monsoon dummy denoted countries subject to the Asian monsoon. An arid dummy variable denoted countries with arid or very dry conditions. A cold dummy variable denoted countries with very cold conditions, such as artic, subartic, or subantartic conditions.

The full analytical dataset included the 2 dependent variables and the 9 independent variables for 32 countries for 1970, 1980, and 1990 and one additional country for 1980 and 1990. The result was 98 observations for both water withdrawals and water consumption. The dataset included three African countries, seven countries from Asia, six from Europe, three from North America, twelve south and Central American countries, and two Oceania countries.

Econometric Model

The objective of the econometric analysis is to estimate an empirical form of equation (12) and test the hypotheses regarding the role of scale, capital per worker, income, and trade openness. The empirical form of equation (12) is specified as

$$w_{it} = x_{it}\gamma + c_i + u_{it} \qquad (13)$$

where w_{it} represents either water withdrawals or consumption per capita, x_{it} is the vector of explanatory variables, γ is a vector of coefficients to be estimated, c_i represents unobserved country effects, u_{it} denotes idiosyncratic disturbances that vary across time and country, and the subscripts $i = \{1,..,98\}$ represent individual countries and the subscripts $t = \{1,..,T_i\}$ represent years.

A troubling problem in country-level analysis is accounting for all the relevant variables that affect the dependent variables while preserving the degrees of freedom required for estimation and hypothesis testing. It is always possible that potentially important variables may be omitted from the analysis. There are two econometric procedures that deal explicitly with the potential impact of omitted variables on estimation and testing with panel data. These are the fixed effects (FE) and random effects (RE) estimators.

A FE estimator treats the unobserved country effects, c_i as constant over time, but varying across countries. A FE estimator deals with one of the problems potentially introduced by omitted variables, correlation between the unobserved effect and the explanatory variables. The FE estimator is asymptotically consistent in the presence of such correlation.

The RE estimator treats the country-level effect as a stochastic variable. The stochastic specification allows for omitted variables that may be constant over time but differ between countries, and those that may be fixed between countries but vary over time. In contrast to the FE estimator, the RE model is consistent when the unobserved effect is uncorrelated with the observable explanatory. The stochastic unobserved effect and the idiosyncratic disturbance term may then be added to form a composite error term, $v_{it} = c_i + u_{it}$. Because c_i appears in the composite error for every time period t, the composite error term is serially correlated. The RE model is estimated using generalized least squares to address serial correlation (Wooldridge 2002).

Equation (13) was estimated for both water withdrawals and water consumption in three stages. At the first stage, equations for water withdrawals and consumption were estimated using a RE estimator and tests were conducted for heteroskedasticity in the data using a likelihood ratio (LR) test. Results rejected homoskedastic errors for the RE estimates at conventional levels of significance. At the second stage, water withdrawal and consumption equations were estimated using a standard FE estimator with robust standard errors and an RE estimator allowing for heteroskedastic errors. To allow estimation of the FE, the FE and RE estimates were estimated without the dummy variables for Asian monsoon, arid conditions, and cold climate. The latter variables could not be included in the FE equations since each dummy is constant across time and the drops out of the FE within estimator.

The RE estimates were evaluated for consistency relative to the FE estimates using a Hausman (1978) test. The Hausman test evaluated the null hypothesis that the estimated coefficients of the random effects model were the same as the fixed effects model estimates. The test indicated no significant difference between the RE and FE estimates at any conventional level of significance, meaning that any correlation between the country level effect and the independent variables was not statistically significant. Hence, attention was focused on the RE estimator and estimates.

The final stage estimated two forms of water use equations. The first form followed the conventional Kuznets curve approach and used a single function quadratic in income to capture the summary effect of the economy on water use,

$$w_{it} = \gamma_0 + \gamma_I Income_{it} + \gamma_{I^2} Income_{it}^2 + \sum_{g=1}^{5} \gamma_{\omega^g} \omega_{it}^g + v_{it} \tag{14}$$

where the $Income_{it}$ is the ith country's gross national income at time t, and ω_{it}^g are the climatic variables discussed in the last section. In contrast to the Kuznets relationship, equation (12) identified four specific and measurable factors that influence water use: scale, composition, income-technique, and openness to trade. Equation (12) therefore leads to an alternative empirical relationship,

$$\begin{aligned} w_{it} &= \gamma_0 + \gamma_S Scale_{it} + \gamma_k Capital_{it} + \lambda_O Openness_{it} + \gamma_I Income_{it} + \gamma_{I^2} Income_{it}^2 \\ &\quad + \sum_{g=1}^{5} \gamma_{\omega^g} \omega_{it}^g + v_{it} \end{aligned} \tag{15}$$

where $Scale_{it}$ was specified as gross domestic product, $Capital_{it}$ represented the composition effect with capital per worker, and $Openness_{it}$ was openness to trade variable as discussed in the last section. Equation (15) retained the quadratic in income portion in order to provide a nested comparison with the Kuznets curve equation (14). The final equations were estimated using a random effects estimator that allowed heteroskedasticity with the unbalanced panel data.

Results

Table 1 describes the variables used in the empirical analysis of water use. The dependent variables were water withdrawals, PCww, and water consumption, PCwc. The predetermined variables were those described above.

Table 1 lists two columns of means. The first column of means lists means computed by giving the data for each country an equal weight. Equal country-level weights treat each country as the unit of observation. The second column of means lists means computed by weighting the data for each country by the country's population. The population weighted means give more weight to country level values for India than a country with a smaller population, such as Spain and Japan. The population weighted means provide a closer approximation to the per capita means for the overall population represented by the sample of countries.

Water withdrawals and water consumption are both about 20% greater in the population weighted means than in the equally weighted means. The percentage differences between equally and population weighted means for economic variables are smaller than those for water use, except for Openness to trade which is more than 40% smaller in the population weighted means than in the equally weighted means.

Table 2 lists the estimated coefficient for the water withdrawal and water consumption equations. The table reports coefficients for both the Kuznets Curve (KC) specified in equation (14) and full model derived from the analysis of the two sector trade model and specified in equation (15). In the discussion below, the full model specified in equation (15) is referred to as the Two-Sector Trade Model (TSTM). In Table 2, there are two sets of estimated coefficients using the TSTM, one for water withdrawals and one for water consumption. Table 2 shows that all the estimated coefficients are statistically different from zero at conventional levels of significance, except for the Income squared coefficients in the water withdrawal equations, the Temperature coefficients, and Constant coefficients for each of the four equations.

The KC is nested within the TSTM insofar as it may be derived by setting the Scale, Capital, and Openness coefficients equal to zero in the TSTM. These restrictions are tested with a likelihood ratio test for the pair of equations for water withdrawals and the pair of equations for water consumption. The test statistic for the restricting the Scale, Capital, and Openness coefficients in the water withdrawals TSTM to zero is 26. The test statistic for the restricted water consumption TSTM is 20. Both of the latter test statistics statistics exceed the 99 percent level Chi-squared critical value of 11.3 with 3 degrees of freedom. Hence, the Kuznets Curve specifications for water withdrawals and water consumption are rejected in favor of the TSTM specifications. For both withdrawals and consumption, the detailed TSTM specifications that assign specific variables to the scale, composition, income-technique, and openness factors explain the data better than the KC specification.

The TSTM coefficients also are consistent with the variable specific hypotheses. The positive coefficients on Scale in the water withdrawal and consumption equations indicate that both forms of water use increase with economic scale. The Scale coefficient for withdrawals is more than twice the size of the coefficient for consumption, so scale has larger impact on withdrawals than on consumption.

The estimated coefficients on Capital are negative and statistically different from zero at the 99 percent level. The estimated Capital coefficient in the water withdrawal TSTM is about 4 times more negative that that the water consumption TSTM. The negative signs are consistent with the hypothesis that capital investment reduces water use by either substituting capital investment for water in the agricultural sector or indicating the comparative advantage of the manufacturing and commercial sectors relative to agriculture.

The estimated Openness to trade coefficients are also statistically different from zero tatht the 99 percent level in both the water withdrawal and consumption TSTM equations. The negative signs on the Openness coefficients are consistent with the hypotheses that openness allows water using activities to move from countries with high water costs to countries where their water costs are lowest. Openness to trade may also facilitate the exchange of ideas and processes the reduce water use and induce incentives for water conservation as a way to reduce costs or improve quality.

The estimated Income coefficients are perhaps the most striking difference between the KC and TSTM estimates. The Income coefficients in both KC equations are positive while the estimated Income coefficients in the TSTM equations are negative. The negative KC coefficients summarize the net effect of the four economic factors, albeit in a statistically imprecise manner. In contrast, the TSTM partitions the four effects to separate variables—Scale, Capital, Openness to trade, and Income. Consistent with the income-technique hypothesis, the signs of the TSTM Income coefficients are negative. Moreover, the Income squared coefficients are also negative in the TSTM, meaning that the water conserving effect of income becomes increasing strong as national income rises. Apparently, increases in income do induce improved water regulation and allocation institutions, leading to improved water using production techniques.

The climate variables have impacts on water withdrawals and consumption that are consistent with intuition. The negative and statistically significant signs on the coefficients for Precipitation indicate that precipitation is a substitute for human water withdrawals and consumption—the more rainfall a country enjoys the less water it needs to withdraw and distribute through human water infrastructure. The signs and significance of Arid are also consistent with the later interpretation. Arid countries lack the natural distribution of water through rainfall and therefore withdraw, distribute, and consume more water than non-arid countries. Nevertheless, the Asian monsoon indicates that water withdrawals and consumption is greater in countries with a seasonal abundance of water from monsoon rains. The Cold dummy coefficients indicate that water withdrawals and consumption are less in countries with short growing seasons and less evaporation.

Finally, the coefficients for Temperature are not significantly different from zero once the other climatic factors are taken into account. Interestingly, these results for Temperature and other climate factors suggest that the effects of global warming on water use are not likely to arise directly through changes in temperature, but indirectly through changes in precipitation, the extent of aridity, and the length of growing seasons in very cold regions of the world.

The estimated TSTM equations were use to simulate water use changes as economies grow. The TSTM equations were used to simulate the impacts on water withdrawals and water consumption of a 10 percent change in the economic variables from their 1990 levels. A 10 percent change Scale and Income is equivalent to a decade of growth for an economy growing at 1 percent per year and 3 years of growth for an economy growing at just over 3 percent. In terms of Openness, a 10 percent change from the Population Weighted mean Openness (Table 1) is less than a 3% increase in the value of exports and imports relative to overall gross domestic product.

Table 3 lists the water use impacts of a 10 percent change in the economic factors of Scale, Capital, Income, and Openness. The results show that water use is highly influenced by economic factors, especially Scale and Income. The 10 percent increase in Scale alone leads to a simulated increase in water withdrawals and consumption of more the 60 percent. However, the economic pressures exerted by Scale are offset by the countervailing pressures of Capital, Income, and Openness. Investment resulting in 10 percent more capital per worker reduces water withdrawals by almost 18 percent and water consumption by almost 10 percent. A 10 percent increase in Income reduces water withdrawals by 44 percent and water consumption by almost 50 percent.

A real economy is unlikely to experience a 10 percent increase in Scale or Capital, without an increase in Income as well. Hence, it is policy relevant to consider the combined effects of changes in Scale, Capital and Income. Table 3 shows that a 10 percent change in Scale, Capital and Income increase mean water withdrawals by 5.4 percent and mean water consumption by 5.2 percent. The latter estimates are similar to those of the International Hydrological Union which forecasts a change of 4.9 percent for water withdrawals and a change of 5.2 percent in water consumption between 1990 and 2000 (IHP, 2005).

Openness to trade also needs to be accounted for in the simulation results. During the decade of the 1990s, there was a significant increase in the amount of world trade (WTO, 2004) and openness to trade is the fourth economic factor with significant impacts on water use. It is also a value that is for the most part subject to governmental control through tariffs and trade policy. Hence, it is of key policy interest.

Interestingly, the simulation results in Table 3 show that a 10 percent change in Openness to trade reduces water withdrawals by almost 9 percent and water consumption by more than 10 percent from their 1990 levels. When combined with a 10 percent change in Scale, Capital and Income, a 10 percent change in Openness results in a net 3.3 percent reduction in water withdrawals and a net 5.3 percent reduction in water consumption. Hence, trade liberalization appears to be a policy level that can shift economic incentives from greater water withdrawals and consumption to greater water conservation.

Conclusions

The analysis demonstrates that national economic structure has profound consequences for global water use. It is not enough to know whether a country is high or low income, is cold or hot, or has more or less precipitation. Economic organization and incentives have a significant impact on water use. Water withdrawals and water consumption vary with economic scale, composition, national income, and the degree of openness to trade in the world economy. Increases in economic scale unequivocally increase water use, but scale is offset by changes in the other economic factors. Increases in capital investment, improvements in national income, and national policies that encourage openness to trade have important effects on reducing water use and improving water conservation.

Results indicate that openness to trade is a key policy variable in reducing water use and improving water conservation. Holding openness to trade constant, mean water use increases by 5 percent after a decade of economic growth of 1 percent per year in scale, capital, and income. However, a 10 percent improvement in openness to trade converts the increase of 5 percent into a reduction of 3 to 5 percent. These results for trade openness reinforce and generalize the sector and commodity specific results of De Fraiture et al (2004). Water conservation could be further improved by policies that encourage capital investment, especially in water saving technologies.

The results show that an understanding of economic structure and organization is essential to improving water use forecasts and policy. The poor performance of forecasts noted by Gleick (2000) is not surprising given the substantial impacts on water use by scale, capital investment, income, and openness to trade. Changes in economic factors, such as capital investment, and changes in economic policy, such as trade liberalization, work to reduce water use and to belie even the best forecasts when such forecasts rely only on past trends in water use. Past water use depends on the past economic conditions, conditions that change as economies develop and the world economy becomes more integrated.

The same economic and policy variables that make water use trend analysis error prone also hold out promise for reducing human water withdrawals and water consumption. Capital investment can substitute improved processes and technology for increased quantities of water. Continued reductions in trade barriers encourage water conservation by allowing the movement of water intensive production activities to countries with lower water costs and environmental conditions that facilitate water conservation. Both capital investment and openness to trade have the potential to shift the global water balance from continuing increases in water use to continuing reductions and improvements in water conservation.

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Table 1. Variable Names, Means, and Descriptions

Variable Name	Variable Means[a]		Description
	Equally Weighted Country Data	Population Weighted Country Data	
PCww	710 (489)	857 (593)	Per capita water withdrawals, cubic meters per year
PCwc	349 (232)	422 (179)	Per capita water consumption, cubic meters per year
Scale	8.52 (7.20)	8.50 (9.04)	Per capita gross domestic product in $US 1,000, 2005 price level
Capital	20.0 (17.4)	17.5 (19.5)	Capital per worker in $US 1,000, 2005 price level
Income	8.33 (7.16)	8.46 (9.09)	Per capita gross national income measure in $US 1,000, 2005 price level.
Income squared	120 (181)	153 (245)	Per capital income squared
Openness	.461 (.209)	.267 (.165)	The sum of exports and imports divided by gross domestic product.
Temperature	292 (6.96)	291 (6.49)	Mean annual temperature measured in degrees Kelvin
Precipitation	929 (476)	1033 (372)	Mean annual precipitation measured in millimeters.
Monsoon	.122 (329)	.527 (.502)	A value of 1 if a country is subject to an Asian monsoon climate; 0 otherwise.
Arid	.306 (.463)	.275 (.449)	A 1 if a country's standard climate description indicates an arid or very dry climate; 0 otherwise.
Cold	.153 (.362)	.109 (.313)	A 1 if a country's standard climate description indicates a very cold climate (e.g., artic or subartic); 0 otherwise.

a. Standard deviations are given in parentheses.

Table 2. Water Use Equation Estimates[a]

Variable	Water Withdrawals[b]		Water Consumption[c]	
	Kuznets Curve	Two-Sector Trade Model	Kuznets Curve	Two-Sector Trade Model
Scale	-	413**		170**
		(97.5)		(56.7)
Capital	-	-36.5**		-8.90**
		(4.30)		(2.31)
Openness	-	-520**	-	291**
		(86.3)		(42.9)
Income	51.3**	-277**	24.0**	-116**
	(12.4)	(94.4)	(6.00)	(54.2)
Income squared	-.296	-.264	-.455**	-.755**
	(.558)	(468)	(.215)	(.187)
Precipitation	-.213**	-.198**	-.166**	-.130**
	(.054)	(.047)	(.031	(.028)
Arid	393**	115**	172**	80.6**
	(42.1)	(51.4)	(30.4)	(29.3)
Monsoon	414**	186**	293**	236**
	(39.7)	(45.8)	(27.7)	(28.9)
Cold	-134**	-258**	-135**	-183**
	(116)	(96.0)	(54.0)	(42.9)
Temperature	-13.6	.914	-1.47	2.04
	(4.85)	(3.69)	(2.24)	(1.31)
Constant	4298	310	704	275
	(1382)	(1048)	(636)	(378)
Log-likelihood	-637	-624	-567	-557
Number of observations	98	98	98	98

a. The standard errors are given in parentheses. A “*” indicates that a coefficient is statistically different from zero at the 95% level. A “**” indicates that a coefficient is statistically different from zero at the 99% level.
b. The dependent variable is PCww.
c. The dependent variable is PCwc.

Table 3. Water Use Impacts of a 10% Increase in the 1990 Economic Variables

Independent Variable Changing by 10%	%Change in Water Withdrawal[a]		%Change in Water Consumption[a]	
	Mean	Median	Mean	Median
Scale	66.9	58.0	62.4	50.6
	(43.3)		(51.3)	
Capital	-17.5	-13.0	-9.2	-6.0
	(20.5)		(11.4)	
Income	-44.0	-37.8	-48.1	-37.0
	(28.0)		(41.2)	
Openness	-8.7	-3.9	-10.5	-5.5
	(11.9)		(14.8)	
Scale, Capital, and	5.4	6.6	5.2	4.0
Income	(7.8)		(7.3)	
			15.2	
Scale, Capital, Income,	-3.3	1.4	-5.3	-1.24
and Openness	(15.6)		(11.4)	

a. The standard errors for the estimated elasticities are given in parentheses below the coefficient estimates.

Citations in text but not in EndNote format

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(International Hydrological Programme (IHP) 2005)
(Grossman and Krueger 1991)
(Barbier 2004)
(De Fraiture et al. 2004)
(Copeland and Taylor 2003)
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Did the Northwest Forest Plan Impact Rural Counties More than Urban Counties?

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Did the Northwest Forest Plan Impact Rural Counties More Than Urban Counties?

Abstract: Controversies over conserving land to protect biodiversity often stem from conflicting views on how conservation will affect local economies. A particular concern is for especially vulnerable rural counties, with low economic diversity and often high proportions of public land. The traditional view is that conserving public land for biodiversity will reduce inputs to commodity production, resulting in local employment declines and out-migration, especially in rural areas. Recently, others have argued that conserving public lands may stimulate local economies by producing amenities that attract firms, workers, and migrants. We study the impacts of the Northwest Forest Plan (NWFP), a massive experiment in reallocating federally-owned land from timber production to ecosystem management, on county employment growth and net migration rates. Our econometric specification allows for different impacts, depending on whether the county is classified as metropolitan, intermediate, or rural. Our results indicate that both metropolitan and rural counties with higher proportions of land allocated to biodiversity protection experienced higher rates of employment growth. Rural counties with higher proportions of protected land also experienced higher rates of net migration, while rural net migration is negatively correlated with higher proportions of land available for timber harvesting. We also find evidence that the economic impacts of the NWFP were short-lived, with all significant impacts occurring while the NWFP was being formulated and disappearing in the decade after its adoption.

Keywords: biodiversity, land conservation, employment, migration, public land, rural counties.

Did the Northwest Forest Plan Impact Rural Counties More Than Urban Counties?

1. Introduction

The 1994 Northwest Forest Plan (NWFP) reallocated 11.5 million acres of federal land in the western Oregon, western Washington and northern California from primarily commodity production, especially timber, to ecosystem management, focused on "maintaining viable populations of native species, native ecosystem types, and evolutionary and ecological processes over long time horizons, accommodating human use and occupancy within these constraints (Charnley 2006a, p. 331)." While federal harvests averaged 4.5 billion board feet (BBF) in the 1980's, the legal and administrative conflicts leading to the NWFP were accompanied by harvests declining to 2.4 BBF between 1990 and 1992 and to 0.5 BBF thereafter. Private harvests did not increase in response to reductions in federal harvests. In fact, they declined by 25% during the 1990s.

All observers predicted negative economic impacts from the NWFP (Charnley 2006a), with job loss predictions ranging from 13,000 (Anderson and Olsen 1991) to 147,000 (Beuter 1990). Most observers predicted that rural counties would be most heavily hit, especially "timber dependent" counties. None of the published analyses predicted or even allowed for potential positive economic impacts from the NWFP (Niemi et al. 1999; Marcot and Thomas 1997). Years after the NWFP was approved, there is still concern that the economic losses (especially jobs) have been heavy and that the losses have been disproportionately borne by rural communities (Spencer 1999, National Research Council 2000, Charnley 2006 a,b,c , Phillips 2006).

Yet, other research challenges the view the decreasing commodity production negatively impacts local economies. Instead, several authors have argued that conservation uses of public lands, rather than commodity production, may actually improve local economies (Power 1996, Duffy-Deno 1998, Lorah 2000, Niemi et al 1999, Power and Barrett 2001, Charnley 2006a, Power 2006). This process may work in any of four ways: 1) conservation lands may attract firms, whose employees value the resulting amenities[1]; 2) conservation lands may provide production inputs for recreation and other natural amenity-based enterprises (Marcouiller and Deller, 1996); 3) firms may be attracted to a pool of workers, who, by migrating, have expressed a willingness to trade income for amenities from conservation lands; and 4) environmental amenities may attract new residents with external sources of income (Lorah 1999; Charnley 2006a).

These conflicting views motivate an empirical investigation of the effects of public land management on local economies. Sufficient time has elapsed to permit us to analyze the impacts of the NWFP on county employment and migration. At least two retrospective studies of the NWFP have already been done. Phillips (2006) estimates actual losses of 30,000 timber industry jobs and 45,000 jobs overall, but does not attempt to evaluate amenity-induced changes in employment. Berck et al. (2003) find that timber employment reductions attributed to the NWFP had little effect on other employment or local poverty in northern California counties.

Evidence is also available from studies of other conservation lands. Duffy-Deno (1998) examines the effect of wilderness areas and finds no impact on employment or population density growth in 250 rural counties. Lorah (2000) finds a positive correlation

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1 Johnson and Rasker (1993) and Crompton et al. (1997) find survey evidence that some firms base relocation decisions more on “quality of life” considerations, including the amenities provided by public land, than on business-related factors.

between 1969-1999 employment, income, and population growth and land in national parks, wilderness, national monuments, and wilderness study areas. Lewis et al. (2002, 2003) find that conservation land designation had no effect on population, employment, or wage growth in the Northern Forest region.

There are four potential reasons why many previous studies have failed to find economic impacts from conservation lands. First, the lands examined may be relatively poor in commodity-oriented resources (Duffy-Deno 1998). Second, conservation lands may have been designated so long ago that all economic adjustments have occurred (Lewis et al. 2002, 2003; Duffy-Deno 1998). Third, the amount of conservation land may be too small to produce measurable effects. Fourth, empirical specifications imposing homogeneous economic effects across heterogeneous communities may mask adverse economic impacts on communities strongly dependent on commodity flows.

This study seeks to surmount all four potential problems. First, much of the re-classified land is among the most productive timberland on the planet and, and second, the changes occurred recently. Third, the area placed under ecosystem management is large, nearly equal to the combined areas of New Hampshire and Vermont. For the counties examined, an average 12% of total land area was reserved from timber cutting. In 11 counties, more than 20% of the land was reserved. In addition, prior studies estimate impacts occurring over ten years or more.[2] We examine shorter periods, delineated to correspond to actual political, scientific, and economic events surrounding the NWFP's development and implementation.

2 For example, Duffy-Deno (1998) and Lorah (2000) analyze the impacts of wilderness areas for 1980-1990 and 1969-1999, respectively, while most wilderness lands were designated in 1964.

The fourth potential reason for failure to find adverse economic effects of conserving land may strongly apply to the NWFP. In the Pacific Northwest (PNW), economic prosperity followed the adoption of the NWFP, as incomes, employment, and populations grew faster than most other regions in the years following the NWFP (Goodstein 1999; Charnley 2006b). But this may have been due to an unrelated stroke of luck in the form of a boom in technology industries in Portland and Seattle and urban growth in these and a few other lucky areas. In other words, did a booming 1990's economy leave the appearance of regional prosperity, with no apparent adverse NWFP impacts, while more isolated, rural counties in fact suffered and are still suffering?

In this paper, we analyze the effects of large-scale biodiversity protection on PNW county employment growth and net migration rates. We examine whether and to what extent the amenities generated by land conservation positively affect local economies and offset negative effects from reductions in commodity-oriented uses of public land. To mitigate the possibility that regional prosperity masks local suffering, we distinguish between impacts on metropolitan, intermediate metropolitan, and rural counties.

We follow previous studies in the migration literature and analyze the simultaneous relationship between county-level employment growth and net migration rates. Our data covers 73 counties containing lands reclassified under the NWFP or adjacent to such counties. In the next section, we briefly review the history of the NWFP. This history guides our econometric analysis and the interpretation of empirical results. Section 3 presents the econometric design and implementation strategy and

Section 4 presents the results. We end with a summary and discussion.

2. A Brief History of the NWFP

Following WWII, federally owned PNW land was managed primarily for timber production. With annual harvests averaging 4.5 BBF in the 1980's, large-scale conversion of old-growth forests to even-aged stands occurred, until only 13 percent of the PNW's old-growth forests remained (Anderson and Olsen 1991).3 Initially, there was little scientific or public concern with the ecological implications of timber harvests. Subsequently, evidence accumulated that old-growth conversion to intensively managed tree farms threatened a variety of ecosystem functions and services, including biodiversity, energy flows, and nutrient and water cycling (Franklin et al. 1981; Forsman et al. 1977; National Research Council 2000). Much of this concern became focused on the northern spotted owl (NSO).

In 1975, Oregon listed the NSO as "threatened", but in 1981 the FWS concluded that threats did not yet warrant an ESA listing (Noon and McKelvey 1996). In 1982, the U.S. Forest Services' (FS) regional planning guide designated the NSO as an indicator species of the region's old-growth forest and provided management guidelines and habitat objectives to preserve the species, including special management areas surrounding 375 known owl pairs. In 1987, the U.S. Fish and Wildlife Service (FWS) conducted a second status review and again concluded that ESA listing was unwarranted. The decision was appealed and, in 1988, Federal District Court ruled that FWS's decision was "arbitrary and capricious" and ordered third review (Noon and McKelvey 1996, p.

3 Pre-European PNW forests were probably composed of 60-70 percent old-growth forests (National Research Council 2000).

136). Meanwhile, environmental organizations were challenging proposed FS and BLM timber sales and, in early 1989, secured a court injunction against the sale of timber from old-growth forests on Bureau of Land Management's (BLM) Oregon land. Federal District Judge Dwyer also issued an injunction against FS timber sales, citing violations of the National Forest Management Act (NFMA) and the National Environmental Policy Act (NEPA) (Marcot and Thomas 1997, p. 3 and Table 1). In June 1989 FWS proposed placing the NSO on the ESA's threatened list.

To side step the ongoing litigation and stabilize PNW timber supplies, U.S. Congress passed a compromise amendment to the appropriations bill funding FS and BLM for fiscal 1990. This so-called Section 318 declared the existing FS Environmental Impact Statement (EIS) and BLM's management plans to be sufficient for preparing PNW timber sales for fiscal year 1990 (Marcot and Thomas 1997, p. 3). Section 318 also directed FS and BLM to delineate "ecologically significant" old-growth timber stands for interim protection while management plans were being prepared (Marcot and Thomas 1997, p. 3) and created the Interagency Scientific Committee (ISC) to address the Conservation of the Northern Spotted Owl. The ISC's mission was to "develop a scientifically credible conservation strategy" for the NSO (Interagency Scientific Committee 1991 p. 24). The May 1990, ISC report concluded that FS and BLM owl management guidelines were a "prescription for [the bird's] extinction." As remedy, ISC proposed 193 Habitat Conservation Areas (HCAs) embedded in a reserve/matrix design spreading HCAs to enable successful dispersal of juvenile owls (Thomas et al. 2006). The goal was to increase the 925 known NSO pairs on proposed HCA's to 1750 by the year 2100 (Wood 1991, p. 40). The HCAs ranged from 50 to 676,000 acres (Thomas et al. 1990), were explicitly mapped, and totaled 7.7 million acres. Nearly half of the acres, however, were already in wilderness areas, national parks, or otherwise unavailable for commercial harvest (Interagency Scientific Committee 1991 p. 24). In June 1990, the FWS listed the NSO as threatened throughout its range in Oregon, Washington, and northern California.

Time magazine featured the controversy in a cover story titled “Owls versus Man” (June 25, 1990). The ESA listing was linked to a putative one-third reduction in PNW federal timber harvests, the loss of 30,000 jobs over ten years, and catastrophe for some communities, especially those in rural areas. The story also aptly described some old-growth forest values in jeopardy if harvests continued (Gup 1990). In September 1990, Assistant Secretary of Agriculture James Mosely issued a press release stating that FS would not adopt the ISC’s recommendations, but would conduct harvests “in a manner not inconsistent with” ISC guidelines. In addition, BLM would follow management guidelines that allowed more harvests than the ISC plan on its Oregon lands (Marcot and Thomas 1997, p. 7).4

Federal District Judge Dwyer enjoined FS timber sales in NSO habitat in May 1991, citing FS’s failure to adopt ISC guidelines. Dwyer ordered FS to produce an EIS consistent with the mandates of NEPA and the ESA (Marcot and Thomas 1997, p. 4). The courts also ordered FWS to propose critical habitat for the NSO (Marcot and Thomas 1997, p. 5). Concurrently, the U.S. House of Representatives chartered the Scientific Panel on Late-Successional Forest Ecosystems (Panel) to assess the viability and integrity of all vertebrate species closely associated with late-successional forests, at-risk fish

4 ISC soon recognized that legal decisions and ESA-related activities of FWS would

stocks, and late-successional forest ecosystems within the NSO's range (Marcot and Thomas 1997, p. 4). The charter was a major turning point because it expanded the officially recognized scope of the issue from the NSO to the viability of other species associated with late-successional and old-growth (LSOG) forests and the integrity of the LSOG ecosystem (Marcot and Thomas 1997, p. 4).

As the Panel deliberated, Seattle Audubon Society, citing the ISC, brought suit against FS, charging that existing guidelines contained in its proposed EIS violated NEPA and NFMA mandates (Marcot and Thomas 1997, p. 4). Portland Audubon Society filed a similar suit against BLM in February 1992. The Panel's October 1991 report concluded that current harvest levels and species viability were incompatible and that establishment of LSOG reserves was the best way to insure the long-term viability of LSOG ecosystems and their component species, including the NSO, marbled murrelet, and anadromous fish (Marcot and Thomas 1997, p. 5). It recommended "owl addition areas" and the protection of "key watersheds." The Panel mapped and classified LSOG forests on federal land and delineated "key watersheds." It presented 14 management alternatives, many of which considered variations in the management of matrix (unreserved) lands between LSOG reserves, owl addition areas, and key watersheds (Thomas et al. 2006, Table 1). It did not recommend a specific alternative (Marcot and Thomas 1997, p. 5).

An ESA "God Squad" Committee, convened by Interior Secretary Lujan, opened 1992 with January hearings in Portland, Oregon. The Committee directed BLM to immediately adopt a still-draft FWS Northern Spotted Owl Recovery Plan. This plan

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require a more draconian strategy than it proposed (ISC 1991, p. 26).

closely followed the ISC strategy, which BLM had previously rejected (Marcot and Thomas 1997, p. 7). In May 1992, Judge Dwyer ruled in favor of Seattle Audubon Society that the FS's EIS violated NEPA guidelines and enjoined FS timber sales in the NSO's range. FS and FWS established a Scientific Analysis Team (SAT) to assess all LSOG vertebrate species, including at-risk fish stocks. SAT then requested and received permission to extend their assessment to all macroorganisms. The published SAT report listed 667 LSOG species and represented a significant step toward a broad ecosystem approach to forest management (Marcot and Thomas 1997, p.10). FWS made a final NSO critical habitat designation consisting entirely of federal land and added the marbled murrelet to its list of threatened species.

As promised during the 1991 Presidential campaign, newly-elected President Clinton convened a televised Forest Conference in Portland, Oregon in April 1993 and ordered a Forest Ecosystem Management Assessment Team (FEMAT) to develop recommendations for managing federal NSO land. In 90 days, FEMAT delivered a report containing a set of ten options for management (Marcot and Thomas 1997, p.11). FS shortly produced a new EIS based on Option 9, the President's choice. Option 9, and the new EIS, designated reserves to preserve LSOG forests and riparian environments, forest matrix lands for timber production, and ten adaptive management areas for testing new silvicultural methods.

A slightly modified Option 9 became FS and BLM policy with the April 1994 record of decision (ROD), signed by the secretaries of Interior and Agriculture. ROD amended the planning documents of 19 National Forests and seven BLM districts. As it came to be known, the NWFP reserved 77 percent of federal land in the NSO's range, with the remainder available for active forest management and/or harvest. It provided for NSO and marbled murrelet habitat, an aquatic conservation strategy to protect at-risk species, including anadromous fish, “survey and manage” guidelines to provide information and protection for lesser-known and potentially vulnerable species of animals, plants, lichen, and fungi (Marcot and Thomas 1997, p.11), and called for National Forest level watershed analysis (Marcot and Thomas 1997, p.12). In addition, the ROD mandated additional harvest restrictions on matrix land (see Noon and McKelvey 1996, p. 158) and projected timber harvests of 1 billion BF per year, a 73 percent reduction from 1980 levels. The strategy was immediately challenged in the courts, but in December 1994 Judge Dwyer ruled that the Option 9 strategy complied with NEPA and the NFMA, simultaneously resolving four ongoing litigations against FS and BLM (Marcot and Thomas 1997, p.17).

The resolution of the conflict, as it culminated in the 1994 ROD, progressed along two fronts. First, the focus shifted from addressing only the habitat requirements of the NSO to the design of forest policy based on a broader concept of ecosystem management. Attendant to this shift were nearly certain prospects of decreasing timber harvests and an increasing fraction of federal forests managed to provide ecosystem services and natural amenities. Initially, public debate (Gup 1990) and the ISC’s guidelines focused on the NSO. However, even the ISC was aware that the NSO was not the real issue and the Panel report of October 1991 addressed the viability of the marbled murrelet and anadromous fish stocks. The 1992 SAT report listed 667 LSOG species and coincided with Judge Dwyer’s decision that the FS’ EIS violated NEPA by not considering other LSOG species. By the end of the period, the ROD codified the design and management of LSOG reserves using principles of ecosystem management and the secondary role of matrix lands, where harvest decisions could be trumped by "survey and manage" guidelines and other requirements.

On the second front, the shape of the eventual resolution became increasingly clear. Initially, environmentalists, scientists, agencies, and industry drew disparate pictures of what public forests should or would look like and communities faced great uncertainty (Gup 1990). Over time, however, the scientific community converged on the principles of extensive LSOG reserves, riparian areas protection, and an even more cautious approach to harvesting on matrix lands. It also became increasingly clear that court approval of FS and BLM actions would require this approach. Although FEMAT's 1993 report contained ten options, President Clinton quickly resolved the remaining uncertainty by choosing Option 9. The period ended with the April 1994 ROD being largely accepted by the courts. Dwyer's December 1994 decision *"…[F]or the first time described the standard of adequate adherence to NFMA regulations, as well as for BLM forest management, pertaining to evaluating and managing for native and desired nonnative vertebrate species and invertebrate species on these public forest lands. It marked for the first time in several years that owl habitats were to be managed by FS and BLM under a common ecosystem management plan found lawful by the courts."* (Marcot and Thomas 1997, p.17).

The events leading to the current NWFP can be grouped into four periods: 1980-1990, 1990-1992, 1992-1994, and 1994-2003. The first period was one of large and relatively stable timber harvests. In the second period federal agencies responded to a changing legal and political climate. Uncertainty prevailed about the type and degree of policy response required and the focus was largely on satisfying legal requirements related to the NSO. However, by the 1992-1994 period, the focus had shifted to include hundreds of species associated with LSOG forests and the LSOG ecosystem itself. The scientific community converged on the need for extensive LSOG reserves, protections for riparian areas, and a cautious approach to harvesting on matrix lands. It became clear that court approval of FS and BLM management plans would require this approach. The final period follows the NWFP's establishment and is long enough to allow the region's economy to adjust to the new policy (Hunt 2006).

In this paper, we empirically examine the latter two periods, 1992-1994 and 1994 –2003. Given the vigorous press coverage, court actions, and the wide-spread engagement of the regions' academic institutions, federal agencies, courts, and government officials, it is reasonable to look for evidence of the NWFP's economic impacts, even though it was not officially adopted until near the end of the period. The second period allows an investigation of the longer-term NWFP impacts and 2003 is the last date for which data are available for many variables.

3. Empirical Modeling Approach

3.1 The Employment–Population Model

We consider the effects of the NWFP on county employment growth (*EG*) and net migration rates (*NM*), two widely used indicators of local economic activity (Hunt 2006). *EG* is measured as the percentage change in county jobs over a specified period. *NM* is the percentage change in county population net of changes due to births and deaths, thus reflecting population changes due to in- and out-migration. In the migration literature, employment and population, or employment change and net migration, are frequently modeled in a simultaneous equations framework (e.g., Greenwood and Hunt 1984, Greenwood et al. 1986, Carlino and Mills 1987, Duffy-Deno 1998, Deller et al. 2001, Lewis et al. 2002, 2003). In growth rate form, the model is[5]:

$$
\begin{aligned}
EG_{jt} &= f_1(NM_{jt}, X_{jt}, \alpha_t) + \varepsilon_t \\
NM_{jt} &= f_1(EG_{jt}, Y_{jt}, \gamma_t) + \mu_t \,, \qquad (1)
\end{aligned}
$$

where EG_{jt} and NM_{jt} are employment growth and net migration rates in county *j* over a specified time period beginning in time *t*, X_{jt} and Y_{jt} are vectors of exogenous variables measured in time *t*, α_t and γ_t are vectors of unknown, time-specific parameters, and ε_t and μ_t are vectors of disturbance terms.

The *EG-NM* model captures the decisions of firms to locate in areas with pools of potential employees and the decisions by individuals to migrate to areas with employment opportunities. The model represents the simultaneous relationship between employment change and migration: migration may induce firms to increase employment in an area while increases in employment may induce further migration. Exogenous variables directly affect EG and NM, but induced changes in *EG* (*NM*) indirectly affect *NM* (*EG*) due to the simultaneity between *EG* and *NM*. If there is a determinant solution to this interactive process, the model can be solved for the total effect (the direct effect

5 Hunt (2006) provides evidence of stationarity in state-level employment and population growth rates. In contrast, levels and log-levels of employment and population are found

plus all indirect effects) of a change in an exogenous variable on *EG* and *NM*.

For the application to the NWFP, we adopt linear specifications of $f_1$and f_2:

$$\begin{aligned} EG_{jt} &= \alpha_{ot} + \alpha_{1t} NM_{jt} + \sum_{k=1}^{3} \alpha_{2kt} Public_{jt} + \alpha_{3t} X_{jt} + \varepsilon_t \\ NM_{jt} &= \beta_{ot} + \beta_{1t} NM_{jt} + \sum_{k=1}^{3} \beta_{2kt} Public_{jt} + \beta_{3t} X_{jt} + \mu_t \ , \end{aligned} \qquad (2)$$

where $PUBLIC_{jt}$ is a vector of variables characterizing public land management in county *j* in time *t* and α_{2kt} and β_{2kt} are the direct effects of public land management on a county of type k in period t.

We allow for heterogeneity in the effects of public land management using the rural Economic Research Service's Rural-Urban Continuum Codes (RUCC) (Economic Research Service 2007). We collapse the ten RUCC into k=three categories, termed METRO, INTER, and RURAL counties based on the counties location vis-à-vis metropolitan areas. METRO counties are those containing a metropolitan area or located on the fringe of an MSA. INTER counties are small, medium and large population counties adjacent to an MSA, and RURAL are small and medium population counties not adjacent to an MSA.[6]

Based on the discussion in section 2, we estimate the model for two distinct periods: 1992-1994, and 1994-2003. EG_{jt} and NM_{jt} are measured as growth rates over these periods and the public land and other exogenous variables are measured as close to the starting year of each period as data availability will allow. All variables are measured

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to be non-stationary in nearly all cases.
6 Specifically, METRO counties are those with RUCC of 0,1,2, or 3, INTER counties are those with RUCC of 8,6, or 4, and RURAL counties are those with RUCC of 9,5,7, or 3. See Cruthers and Vias (2005) for an earlier application of the RUCC.

for 73 counties with federal land reclassified by the NWFP or adjacent to such counties.7 Definitions and notation for the variables are in the Appendix.

3.2 Public Land Management Variables

Table 1 presents the federal land classifications under the NWFP. Only some federally managed land was reclassified. The 7.3 million acres reserved by act of Congress and the 1.5 million acres of administratively withdrawn land were not reclassified and the 4 million acres classified as matrix land was to remain devoted to timber production. Reclassification occurred on the 11.6 million acres in Late Successional Reserves, Managed Late Successional Reserves, Riparian Reserves, and Adaptive Management Areas (see Table 1). In addition, considerable acreages of state-owned forestlands occur across the PNW.

We characterize public land management using five variables. $WILD_{jt}$ is the percentage of county *j*'s land area in wilderness in time t, irrespective of management agency. $NATPARK_{jt \text{ and}}$ $STATE_{jt}$ is the proportion of county *j* land in time t managed by the NPS, and state forestry departments, excluding wilderness lands. Wilderness areas and NPS lands are managed for non-consumptive uses (e.g., recreation), whereas other FS, BLM and state forests were managed for timber production and other uses. Lands reclassified by the NWFP are combined into two categories. $RESERVED_j$ is the proportion of the county's total land area classified as Late Successional Reserves, Managed Late Successional Reserves, or Riparian Reserves. $MATRIX_j$ is the proportion

7 We omit the California counties of Marin, Napa, and Alameda because their economies are likely to be atypical and because they had only tiny NWFP land allocations. The associated adjacent counties of San Mateo, Santa Clara, Contra Costa, and Solano were

classified as Matrix lands or Adaptive Management Areas, since these two classifications were to be the source of 90% of expected harvests (Charnley 2006a).[8]

To account for heterogeneous responses to public land management we interact $RESERVED_j$ and $MATRIX_j$ with the county types derived from the RUCC, to form the NWFP land management variables: $METRO\text{-}RESERVED_j$ (reserved land as a proportion of county land for metro counties, and zero otherwise), $METRO\text{-}MATRIX_j$ (matrix land as a proportion of county land for metro counties, and zero otherwise), $INTER\text{-}RESERVED_j$ (reserved land as a proportion of county land for intermediate counties, and zero otherwise), INTER-MATRIXj (matrix land as a proportion of county land for intermediate counties, and zero otherwise), RURAL-RESERVEDj (reserved land as a proportion of county land for rural counties, and zero otherwise), and $RURAL\text{-}MATRIX_j$ (matrix land as a proportion of county land for rural counties, and zero otherwise). Because the land management variables are measured as shares of *total* county land, they capture effect of diverting land from private uses as well as alternative management objectives. In the *EG* equation, these variables represent production amenities (e.g., a raw material input in the case of timber). In the *NM* equation, these variables represent

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also omitted.

8 The Northwest Forest Plan Regional Ecosystem Office combines the matrix and riparian reserves categories, though lands in these two classifications are managed very differently. Espy and Babbit (1994) report the total area of riparian reserves (Table 1), but none of the available documentation provides a source for this figure. Thus, we calculate county-specific riparian reserve acreage. We use available GIS data to compute the length of rivers and streams found within the matrix/riparian reserves classification. Then, we use NWFP default buffer widths (300 feet for perennial streams and 100 feet for intermittent streams) to compute the area of stream buffers. We find the area of riparian reserves is 1,242,238 acres, about one-half of the figure in Espy and Babbitt (1994). We calculate matrix land acreage by subtracting the calculated acreage of riparian reserves from the area in the combined matrix/riparian reserves category.

consumption amenities to residents and potential migrants.

We use the NWFP land management variables discussed above in the 1992-1994 period because, by the beginning of the period, a clear picture of the policy response had emerged. The Panel's October 1991 report mapped LSOG forests on federal land, identified key watersheds, and presented explicit management alternatives with hundreds of species in mind. The FEMAT report describing the approach taken under the NWFP was not commissioned until April 1993. However, the completion of the report in just 90 days and the near immediate adoption of Option 9 suggest that the NWFP was by this time a *fait accompli*.

To capture potential spillover effects from the NWFP (e.g., consumption amenities and cross-county shipments of harvested timber), we include an indicator variable $ADJNWFP_j$, equal to unity if a county does not contain NWFP land, but is adjacent to one that does. Finally, lands classified under the NWFP as administratively withdrawn are included in the wilderness share, $WILD_{jt}$, as these lands often have characteristics similar to lands in wilderness areas and often are later designated as wilderness.

It is important to note that some public lands in the study region were designated well before the periods we analyze.9 The effects of these designations are probably not reflected in current growth rates, but may be present in employment and population levels. We include lagged employment density, $EMPDEN_{jt}$, defined as total county employment divided by county land area in the EG equation, to control for the historical

9 For example, Mount Rainer and Crater Lake National Parks were established in 1899 and 1902, respectively. Two notable exceptions are Redwoods National Park, expanded in 1978, and small expansions of some wilderness areas in 1984.

effects of conservation lands on employment levels and to isolate their impacts on *EG* during subsequent periods. Similarly, we include lagged population density, POPDEN_{jt} in the *NM* equation.

3.3 Other Exogenous Variables Affecting Employment Growth and Net Migration Rates

The primary goal of the econometric estimation is to obtain consistent and precise estimates of the effects of the NWFP land classifications on *EG* and *NM*. As such, we include many potentially relevant exogenous variables in the model to minimize the potential inconsistency of the parameter estimates. In the case of variables with the potential of being simultaneously determined with EG and NM, we use values for the previous period. In the *EG* equation we include variables to control for factors affecting production costs. Much of the PNW's economic activity is concentrated along the Interstate 5 corridor (Anderson and Olson 1991). We include an indicator variable, INTER5_{jt}, equal to unity if Interstate 5 is located in the j^{th} county. To control for other transportation costs, we include ROADDEN_{jt}, the lagged ratio of interstate and other major arterial road miles to county land area. To control for cost reductions due to spatial proximity to other firms, we include METRO_{jt}, equal to unity if the county is part of a metropolitan statistical area and zero otherwise in the previous period. Because agglomeration effects may be greater in larger urban areas, we include a separate indicator variable, BIGMETRO_{jt}, for counties that are part of either the Portland or Seattle MSAs.

The next set of variables included in the *EG* equation measures factors affecting local labor market conditions. Lagged high school and college graduation rates measure the educational attainment of the local labor force. HSGRAD_{jt} and COGRAD_{jt} are the percentages of the counties' populations over 25 years of age who have completed high school and college, respectively. $EDUCEXP_{jt}$, the lagged share of local government expenditures devoted to education, measures education quality. $UNEMPLOY_{jt}$ is the lagged unemployment rate, and indicates the availability of a local labor force. The direct employment effects of the NWFP may be proportionate to the county' timber dependence, so we include $WOODEARN_{jt}$, equal to lagged total payroll for SIC 24 (lumber and wood products) as a percentage of total county payroll.

Another factor likely to influence a county's employment growth is the log export market. Log exports grew in importance from the 1950's until exports totaled more than 4.5 BBF in 1988 and 1989. Price premiums for export versus domestic logs ranged from 25 to 66 percent, although some of the differential was related to quality (Daniels 2005). The 1990's saw a diminution of the export market arising from a variety of factors, including increasing supply from Canadian forests, changes in stumpage prices, Asian macroeconomic conditions, exchange rates, and Asian preferences for building materials (Daniels 2005).

To reflect the potential importance of the timber exports, we construct a variable, EXPORT, based on the gravity trade model. These models predict bilateral trade flows will vary directly with economic size and inversely with distance. We measure economic size as the percentage of county land held in forest land which could produce logs eligible for export. By 1990 only timber from private land in all three states was eligible for export (see Daniels 2005 for details), so we measure size as the percentage of county land area in private forests. To measure distance, we divide by the square of the distance from the county seat to the nearest seaport suitable for exporting logs (Warren 1989).

Finally, $DIVIDEND_j$ is the lagged share of the county's total personal income derived from dividends and $FEDEXP_j$ is lagged per capita federal expenditures and obligations within the county. These variables control for external income injected into the local economy.

McGranahan's (1999) work suggests that natural amenities (or disamenities) influence *NM*. There are considerable climatic differences in the region, particularly along an east-west gradient. Our climate variables include January and July temperatures ($JANTEMP_j$ and $JULTEMP_j$), hours of sunshine in January ($JANSUN_j$), July humidity ($JULYHUMID_j$), and rainfall in January ($JANRAIN_j$), all computed as means over 1941-1970. We also control for amenity differences by including an indicator variable, $COAST_j$, if the county is adjacent to the Pacific Ocean.

The exogenous variables in the *NM* equation measure the attractiveness of the county to potential migrants and current residents. $INTER5_j$ and $ROADDEN_j$ control for accessibility, which is expected to positively influence migration decisions. $METRO_{jt}$ and $BIGMETRO_j$ may indicate better employment prospects, and/or the attraction of urban amenities. We include all of the land management and climate variables in the *NM* equation to capture the associated amenities.

Additional variables in the *NM* equation control for community characteristics. The expenditure-to-tax ratio ($EXPTAX_{jt}$) is the ratio of local government expenditures to local taxes and proxies for the provision of public sector goods relative to residents' tax burdens. The composition of local government spending is measured by the shares of government expenditures allocated to health care and hospitals, $HEALTHEXP_{jt}$, and education, $EDUCEXP_{jt}$, respectively. To control for the stability of the community, we include the percent of owner-occupied homes, OWNHOME$_{jt}$. Median household income, INCOME$_{jt}$, proxies for several factors, such as the extent and nature of consumer and cultural opportunities and characteristics of the local housing market. Finally, we include CRIME$_{jt}$, measured as the number of serious crimes per 100,000 population.

In both equations, we include separate indicator variables for counties in Oregon, OREGON$_{j}$, and Washington, WASHINGTON$_{j}$, with California counties serving as the reference. These variables control for unmeasured differences between states, such as tax rates, land-use regulations, and state-level public expenditures.

3.4 Estimation Issues

We estimate (2) using Three-Stage Least Squares. Identification of the parameters is achieved through exclusion restrictions. Specifically, from the *EG* equation we exclude most of the community characteristics variables (EXPTAX$_{jt}$, HEALTHEXP$_{jt}$, CRIME$_{jt}$, OWNHOME$_{jt}$, INCOME$_{jt}$) and population density, POPDEN$_{jt}$. From the *NM* equation, we exclude variables for labor market conditions HSGRAD$_{jt}$, COGRAD$_{jt}$, UNEMPLOY$_{jt}$, WOODEARN$_{jt}$, DIVIDEND$_{jt}$, employment density, EMPDEN$_{jt}$, and EXPORT$_{jt}$.

Two estimation issues deserve attention. The first concerns endogeneity of the NWFP variables measuring reserved and matrix land. It was widely perceived that conserving land to protect biodiversity would result in job losses (Niemi et al. 1999) and there may have been political pressure to reserve less land in counties with stagnant economies or high timber dependency. If so, *EG*, *NM*, and the NWFP land management variables may be determined simultaneously. Moreover, the NWFP variables may be subject to measurement error if land classifications were poorly understood or perceived with error by decision makers. Either problem could lead to inconsistent parameter estimates.

We test for biases due to simultaneity or measurement error in the 1992-1994 and 1994-2003 periods using an instrumental variables approach outlined by Geroski (1982). We construct projections of the reserved and unreserved proportions using weighted least squares log-odds regressions:

$$\ln\left(\frac{p_j}{1-p_j}\right) = \delta Z_j + \lambda_j \ , \qquad (3)$$

where p_j is equal to either $RESERVED_j$ or $MATRIX_j$, and Z_j is a vector of instrumental variables (discussed below), δ is a parameter vector, and λ_j is the disturbance term (see Judge et al. 1988, p. 790). We use the estimation results to compute differences between the actual and predicted values of $RESERVED_j$ and $MATRIX_j$. These residuals are then included as regressors in Two-Stage Least Squares estimation of (2). We use *t* and *F* statistics to test the null hypotheses that the parameters on the residuals are separately or jointly zero. If we reject the null with either test, we use fitted values of the variables from (3) in the estimation of (2).

The instruments are the number of NSO centers in the county, the number of marbled murrelet centers, and an indicator variable for whether or not a key watershed is located within the county.[10] These variables reflect ecological criteria used to reclassify

10 Michael Soules kindly provided us with these data. The NSO variable is the number of land-use tracts within a county with at least one NSO center, where centers indicate a single owl or owl pairs. There are 22,869 land-use tracts defined for the region. The marbled murrelet variable is similarly defined. A key watershed is one containing habitat

land under the NWFP and were found by Soules (2002) to be strong predictors of the NWFP allocations. In contrast, the ecological variables are likely to be orthogonal to employment growth and net migration rates. The ecological variables cannot be influenced by local economic conditions (e.g., spotted owls presumably do not consider rates of employment growth when selecting nesting sites). As well, the ecological criteria are weakly related to federal timber supplies and amenities from federal lands and, therefore, unlikely to be determinants of the economic indicators. Soules (2002) found a positive relationship between NSO centers and the classification of unreserved lands. Also, there is evidence that NSOs occupy second- and third-growth stands, which have low timber volumes relative to LSOG stands (Thome et al. 2000). The presence of endangered fish species, as reflected in the watershed variable, does not necessarily indicate large timber volumes on federal lands. Potential migrants are unlikely to be attracted by endangered species *per se*, but rather by amenities whose value is determined by access and other factors.

The second estimation issue is spatial dependence in the residuals (Kelejian and Prucha 2004). Spatial autocorrelation may arise from cross-county effects of the exogenous variables on *EG* and *NM* since we model only within-county effects. We use the residuals from the second stage to estimate the spatial lag parameters, where the elements of the row-standardized weights matrix equal unity for adjacent counties (queen contiguity) and zero otherwise. For each lag parameter, we test the null hypothesis that the estimate is equal to zero. When the null is rejected, we adjust the corresponding residuals for spatial autocorrelation.

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for potentially threatened fish species. These variables are constructed from the actual

4. Results

In the following sub-sections, we discuss the structural estimates in each the two periods. The last subsection presents the total (or solved-structure) effects of the land management variables.

4.1 The 1992-1994 period

We fail to reject the null hypothesis of no spatial autocorrelation for the 1992-1994 period. The instrumental variables tests indicate that the NWFP variables are endogenous in the *EG* equation and exogenous in the NM equation. These projections show that *RESERVED* is positively related to the three ecological instruments and *UNRESERVED* is positively related to the number of NSO centers and the presence of a key watershed, but negatively related to the number of marbled murrelet centers.

Results are presented in Table 2. The R^2 statistics are 0.54 and 0.72 in the *EG* and *NM* equations, respectively. Although the coefficient on *NM* is positive, it is not statistically indistinguishable from zero in the *EG* equation. *EG* appears to be higher in Oregon and Washington, *ceteris paribus*, positively related to the unemployment rate at the beginning of the period, and positively related to sunshine in January (JANSUN). EG is lower in counties more dependent on the wood products industry (WOODEARN), those adjacent to the ocean (COAST), those on the Interstate 5 corridor (INTER5), and with more winter rain (JANRAIN).

In the *NM* equation, the coefficient on *EG* is positive and significantly different from zero. Other variables with positive and significant (5% level) effects on *NM* are the

data used in the NWFP planning process.

Oregon and Washington dummies (*OREGON*, *WASHINGTON*), the home ownership rate (*OWNHOME*), and January sun (JANSUN). Counties with higher federal expenditures (FEDEXP) experienced less NM.

Among the land management variables, the proportions of county land in state forests (*STATE*) and wilderness (WILD) have a positive and significant (10% level) effect on EG. The direct EG effects of reserved land on metro counties (METRORESERVED) and rural counties (RURAL RESERVED) are positive and statistically significant (p-value <=.09). A similar pattern is observed for NM, with positive direct effects for reserved land and negative direct affects for matrix land. However, these coefficients are only statistically significant for RURALRESERVED (p-value<=.07) and for RURALMATRIX (p-value<=.04).

4.2 The 1994-2003 period

For this period, we reject the null hypothesis of spatial independence in the residuals and apply the procedure, discussed above, to adjust the estimates for spatial autocorrelation. We fail to reject the null hypotheses of exogeneity of the NWFP land allocations. The estimation results are in Table 5.

The explanatory power of the model remains relatively high, with R^2 statistics of 0.55 and 0.65 for the *EG* and *NM* equations, respectively. As with the 1992-1994 model, the NM coefficient in the *EG* equation is positive, but statistically insignificant. This may be due to the high proportion of migrants to the region who are retired or otherwise move to the region primarily for its amenities, since these migrants are less likely to enter the labor market directly.[11] As expected, the shares of high school and college graduates

11 A recent survey found that between 40% and 60% of recent migrants to

(*HSGRAD*, *COGRAD*) positively and significantly (5% level) affect employment growth. As in the 1992-1994 period, counties with greater dependency on the wood products industry (*WOODEARN*) experienced significantly lower EG. The results suggest that EG is positively influenced by higher July temperatures and inversely related to lagged employment density. The latter reflects higher EG in more rural counties.

In the *NM* equation, the coefficient on *EG* is positive and significantly different from zero. *NM* is significantly higher in Oregon, Washington, and metropolitan counties (*OREGON*, *WASHINGTON*, *METRO*) and positively related to population density (*POPDEN*), and January temperatures and sun (*JANTEMP*, *JANSUN*), but inversely related to JULYTEMP. *NM* is lower in coastal counties (*COAST*) and counties with higher incomes (*INCOME*), higher federal expenditures (*FEDEXP*), higher road densities (*ROADDEN*), July temperatures (*JULYTEMP*), and summer humidity (JULYHUMID).

Turning to land management effects, we find a significantly positive effect on net migration for state-owned forest land, for reserved land in metro counties (METRO-RESERVED) and intermediate counties (INTER-RESERVED) and for matrix land in rural counties (RURAL MATRIX).

4.3 Total Effects of the Land Management Variables

Each parameter estimate discussed above measures the direct marginal effect of an exogenous variable on either *EG* or *NM*. A change in one endogenous variable, however, affects the other, and so on. By solving the structural equations in (2), we obtain the total effect of an exogenous variable on each endogenous variable. The total effect represents

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Oregon were motivated by amenity considerations only (Judson et al. 1999).

an equilibrium in the sense that it accounts for all adjustments in the endogenous variables. Estimates of total effects for NWFP land management variables and their standard errors (obtained by the delta method) appear in Table 3. As before, we estimate different total effects for each of the three county types.

In the 1992-94 period, we find a significant positive effect of NWFP reserved land on EG for metro counties and for rural counties, while none of the other total EG effects are statistically significant. During the same period, we estimate that NWFP reserved land had a positive NM effect in rural counties, while reserved land appears to have had no NM effect for metro or intermediate counties. Conversely, the NM effect of matrix land for rural counties is significantly negative. For 1994-2003, a pattern of no impacts of the NWFP land classifications on either EG or NM emerges, as none of the estimated total effects are statistically significant at conventional levels.

5. Discussion and Conclusions

Controversies over conserving land for biodiversity protection often stem from differing views on local economic effects. The traditional view is that setting aside public lands for biodiversity will reduce commodity inputs, resulting in lower local employment and out-migration, especially in "timber dependent," rural counties. More recently, others have argued that conservation of public lands may stimulate local economies by producing amenities that attract firms, workers, and migrants.

To date, most empirical evidence suggests that public land conservation has little to no effect on local economic indicators. However, these earlier studies target policies affecting unproductive or small areas of land, or that were implemented long before the period of analysis, and do not allow for heterogeneous economic effects.

In contrast, we study the impacts of the Northwest Forest Plan (NWFP), a massive experiment in reallocating highly productive public forest land from timber production to ecosystem management. We study two time periods, the first period during planning and the second period immediately following the NWFP's implementation. Our empirical strategy also allows us to estimate different effects for metropolitan, intermediate, and rural counties.

The 1992-1994 results suggest that counties with a greater proportion of land reserved for ecosystem management experienced higher employment growth rates after all simultaneity between employment growth and net migration is accounted for. These total impacts are statistically significant for metropolitan and rural counties. We also find that reserved land has a significantly positive effect on net migration for rural counties. Conversely, we find that a larger percentage of rural county land potentially managed for timber production is associated with significantly lower net migration. These results suggest that, rather than being disproportionately hurt by NWFP land protection, rural counties experienced equal or larger economic benefits as the amenities associated with the NWFP led to higher employment growth and net migration and the disamenities associated with commodity production led to less net migration.

Our results also suggest that economic impacts are short lived, even for a policy change as massive as the 11.5 million acre NWFP. In the decade following the implementation of the NWFP, we fail to find evidence of statistically significant effects on county-level employment growth or net migration for rural, intermediate, or metropolitan counties.

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TABLE 1.
NORTHWEST FOREST PLAN LAND CLASSIFICATIONS*

CLASSIFICATION	DEFINITION*	ACRES
CONGRESSIONALLY RESERVED AREAS [1]	RESERVED BY ACT OF CONGRESS, E.G. WILDERNESS AREAS, WILD AND SCENIC RIVERS.	7,320,600
LATE SUCCESSIONAL RESERVES	DEDICATED TO MAINTAINING A FUNCTIONAL, INTERACTIVE, LATE-SUCCESSIONAL AND OLD-GROWTH FOREST. DESIGNED TO SERVE AS HABITAT FOR OLD-GROWTH RELATED SPECIES INCLUDING THE NORTHERN SPOTTED OWL.	7,430,800
ADAPTIVE MANAGEMENT AREAS	DESIGNED TO DEVELOP AND TEST NEW MANAGEMENT APPROACHES TO INTEGRATE AND ACHIEVE ECOLOGICAL, ECONOMIC, AND OTHER SOCIAL AND COMMUNTY GOALS.	1,521,800
ADMINISTRATIVELY WITHDRAWN AREAS	IDENTIFIED IN CURRENT FOREST AND DISTRICT PLANS, OR DRAFT PLAN PREFERRED ALTERNATIVES AND INCLUDE RECREATION AND VISUAL AREAS, BACK COUNTRY, AND OTHER AREAS NOT SCHEDULED FOR TIMBER HARVEST.	1,477,100
MANAGED LATE SUCCESSIONAL RESERVES	EITHER DELINEATED FOR MAPPED, KNOWN SPOTTED OWL ACTIVITY CENTERS OR UNMAPPED PROTECTION BUFFERS, OR DESIGNATED TO PROTECT CERTAIN RARE AND LOCALLY ENDEMIC SPECIES.	102,200
RIPARIAN RESERVES [2]	AREAS ALONG ALL STREAMS, WETLANDS, PONDS, LAKES, AND UNSTABLE AREAS WHERE CONSERVATION OF AQUATIC AND RIPARIAN-DEPENDENT TERRESTRIAL RESOURCES RECEIVES PRIMARY EMPHASIS.	2,627,500
MATRIX	FEDERAL LAND OUTSIDE OF THE SIX CATEGORIES ABOVE. THE AREA IN WHICH MOST TIMBER HARVEST AND OTHER SILVICULTURAL ACTIVITIES WILL BE CONDUCTED. ALSO CONTAINS NON-FORESTED AREAS AND FORESTED AREAS NOT TECHNICALLY SUITED FOR TIMEBER PRODUCTION.	3,975,300

* QUOTED FROM OR A SUMMARY OF ESPY AND BABBITT (1994)

[1] NO NEW LANDS IN THIS CLASSIFICATION WERE ALLOCATED BY THE NORTHWEST FOREST PLAN

[2] SEE THE TEXT FOR FURTHER DISCUSSION ON COMPUTATION OF THE AREA OF RIPARIAN RESERVES

TABLE 2. THREE-STAGE LEAST SQUARES PARAMETER ESTIMATES FOR EMPLOYMENT GROWTH AND NET MIGRATION, 1992-1994*				
	EMPLOYMENT GROWTH		NET MIGRATION	
	PARAMETER ESTIMATE	ABSOLUTE T-STATISTIC	PARAMETER ESTIMATE	ABSOLUTE T-STATISTIC
Constant	-0.383	0.859	-0.258**	3.51
NM	.127	0.17		
EG			0.105**	2.74
OREGON	0.234**	2.96	0.044**	3.81
WASHINGTON	0.205**	2.34	0.051**	3.81
INTER5	-0.078**	3.01	0.002	0.30
METRO	0.050	1.55	0.010	1.46
INCOME			-0.0000004	0.65
HEALTHEXP			0.029	0.86
EDUCEXP	-0.145	1.23	-0.023	0.77
FEDEXP	-0.0045	0.83	-0.002*	1.71
POPDEN			-0.022	1.15
EXPTAX			0.00008	0.06
OWNHOME			0.003**	5.37
ROADDEN	66.90	0.34	6.81	0.15
CRIME			0.000004**	1.98
HSGRAD	0.003	0.77		
COGRAD	0.001	0.45		
WOODEARN	-0.330**	3.407		
DIVIDEND	0.256	1.02		
EMPDEN	-0.101	0.76		
UNEMPLOY	0.024**	3.56		
EXPORT	-0.0000	0.48		
JANTEMP	0.003	1.04	0.0004	0.64
JANSUN	0.001**	2.06	0.0002**	1.92
JULYTEMP	-0.004	1.50	0.0004	0.54
JULYHUMID	0.0007	0.63	0.0003	1.19
JANRAIN	-0.006*	1.75	-0.0007	0.84
BIGMETRO	-0.006	0.19	0.0004	0.05
COAST	-0.090**	2.24	-0.006	0.67
WILD	0.684*	1.91	-0.006	0.08
STATE	0.052*	1.71	0.005	0.71
NATPARK	-0.356	0.30	-0.088	0.31
METRORESERVED	0.309**	2.58	0.0003	0.01
METROMATRIX	-0.075	0.78	-0.014	0.62
INTERRESERVED	0.283	1.04	0.062	1.02
INTERMATRIX	-0.11	0.36	-0.053	0.75

RURALRESERVED	0.455*	1.79	0.091*	1.73
RURALMATRIX	0.479	1.09	-0.217**	2.34
ADJNWFP	0.32	1.08	-0.06	0.95
Mean (S.D) of Dependent Variable	0.063 (0.091)		0.025 (0.026)	
R-Squared	0.544		0.723	

Note: ** indicates significance at the 5% level, * indicates significance at the 10% level

TABLE 3. THREE-STAGE LEAST SQUARES PARAMETER ESTIMATES FOR EMPLOYMENT GROWTH AND NET MIGRATION, 1994-2003*				
	EMPLOYMENT GROWTH		NET MIGRATION	
	PARAMETER ESTIMATE	ABSOLUTE T-STATISTIC	PARAMETER ESTIMATE	ABSOLUTE T-STATISTIC
Constant	-2.165**	-2.91	0.544	1.89
NM	0.470	1.081		
EG			0.447**	5.70
OREGON	0.042	0.33	0.131**	3.28
WASHINGTON	0.007	0.048	0.188**	3.89
INTER5	0.001	0.027	-0.029	-1.53
METRO	-0.049	-0.78	0.065**	2.71
INCOME			-0.000008**	-3.55
HEALTHEXP			0.067	0.568
EDUCEXP	-0.237	-0.87	-0.083	-0.639
FEDEXP	-.002	-0.12	-0.014**	-2.81
POPDEN			0.233**	3.29
EXPTAX			-0.003	-1.46
OWNHOME			0.002	1.12
ROADDEN	521.8454	1.47	-673.86**	-3.99
CRIME			-0.00001	-1.51
HSGRAD	0.019**	2.60		
COGRAD	0.015**	2.74		
WOODEARN	-0.27**	-1.96		
DIVIDEND	0.153	0.392		
EMPDEN	-0.432*	-1.68		
UNEMPLOY	0.006	0.64		
EXPORT	-.00000001	-1.26		
JANTEMP	-0.002	-0.41	0.007**	3.36
JANSUN	-0.0007	-0.54	0.0009**	1.98
JULYTEMP	0.012*	2.18	-0.010**	-4.19
JULYHUMID	0.001	0.52	-0.003**	-2.74
JANRAIN	0.007	1.04	0.001	0.46
BIGMETRO	0.021	0.327	-0.032	-0.99
COAST	-0.014	-0.23	-0.066**	-2.45
WILD	0.111	0.18	0.213	0.81
STATE	-0.060	-1.08	0.052**	2.14
NATPARK	-2.07	-0.90	-0.428	-0.43
METRORESERVED	0.066	0.29	0.211	2.34
METROUNRESERVED	0.245	1.33	-0.125	-1.48
INTERRESERVED	0.562	1.06	-0.518	-2.27
INTERUNRESERVED	-0.80	-1.37	0.722	2.94

RURALRESERVED	0.279	0.66	-0.147	-0.83
RURALUNRESERVED	-1.621	-1.68	1.232	3.14
ADJNWFP	-0.0096	-0.08	-.078	-0.34
Mean (S.D) of Dependent Variable	0.172 (0.167)		0.051 (0.083)	
R-Squared	0.554		0.653	

Note: ** indicates significance at the 5% level, * indicates significance at the 10% level

TABLE 4

TOTAL EFFECTS OF RESERVED AND MATRIX LAND BY COUNTY TYPE*

Variable Name	1992-1994		1994-2003	
	Employment Growth	Net Migration	Employment Growth	Net Migration
METRORESERVED	0.313**	0.033	0.0424	0.231
	(2.64)	(1.04)	(.153)	(1.33)
METROMATRIX	-0.078	-0.022	0.236	-0.020
	(0.78)	(0.79)	(0.96)	(0.13)
INTERRESERVED	0.295	0.093	0.402	-0.339
	(1.06)	(1.22)	(0.59)	(0.80)
INTERMATRIX	-0.121	-0.065	-0.583	0.461
	(0.38)	(0.74)	(0.80)	(1.01)
RURALRESERVED	0.473**	0.141**	0.265	-0.03
	(1.97)	(2.18)	(0.48)	(0.09)
RURALMATRIX	-.513	-.271**	-1.319	-0.643
	(1.21)	(2.37)	(1.10)	(0.86)

Note: ** indicates significance at the 5% level. Absolute values of the t-statistics, computed by the delta method, are in parentheses.

PART 4: Papers Supporting Objective: "Estimate the Economic Value of Agricultural Land Preservation and Open Space"

Willingness to Pay for Land Preservation Across States and Jurisdictional Scale: Implications for Benefit Transfer

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WILLINGNESS TO PAY FOR LAND PRESERVATION ACROSS STATES AND JURISDICTIONAL SCALE: IMPLICATIONS FOR BENEFIT TRANSFER

ABSTRACT

In stated preference valuation of farmland preservation, respondents are often told that preservation will occur within a given jurisdictional scale—e.g., community, state, county—but do not know the specific location of parcels in question. Hence, welfare estimates may be available for different scales and/or states, providing numerous avenues for benefit transfer. This paper provides a systematic assessment of transfer error, contrasting different methods for the transfer of farmland preservation values across states and jurisdictional scales. Empirical results drawn from multi-state choice experiments suggest that the choice of across-scale versus across-state transfer method can have critical implications for transfer validity.

RUNNING TITLE

Land Preservation Benefit Transfer

I. INTRODUCTION

Notwithstanding over two decades of stated preference (SP) research[1] measuring farmland amenity values, little published work addresses the potential transferability of valuation estimates across policy contexts (Bergstrom and Ready 2005). This lack of research is striking, particularly given the relevance of farmland amenity values for policy (Irwin et al. 2003) and the ubiquity of benefit transfer in policy analysis (Bergstrom and De Civita 1999). Nonetheless, with the exception of the literature review of Bergstrom and Ready (2005) and unpublished work of Ozdemir et al. (2004), the authors are aware of no research that provides findings relevant to the transfer of farmland amenity values.

The transferability of farmland amenity values is particularly germane with respect to the issue of scale, defined as the size of the jurisdiction or area over which a given amount of land is preserved. Most SP research occurs at the political-boundary scale since these jurisdictions offer the most realistic funding and implementation mechanisms.[2] For example, community or regional farmland valuation studies (e.g., Bergstrom et al. 1985; Halstead 1984; McLeod et al. 2002) typically estimate willingness to pay (WTP) for delivery of user and nonuser services within an identified locality such as a town or county. In such cases, the scale ensures a close proximity between service delivery and respondents' homes, leading to the potential for significant use and nonuse values. In contrast, statewide studies (Duke and Ilvento 2004; Ozdemir et al. 2004) generally solicit WTP for preservation at the state scale, with a concomitant expectation that preserved land will not be located close to respondents' homes. Given that proximity to preserved farmland is not expected at the state scale, nonuse values become the primary motivation for survey responses (although limited use values might also be anticipated).

The issue of scale also is relevant for farmland preservation because funding decisions, such as referenda on preservation bonds, are typically made before the identities of targeted parcels are known. SP surveys replicate this lack of spatial certainty and rarely specify the exact location of targeted farmland parcels. Respondents are simply told that farmland preservation will occur somewhere within a given community, county, or state—thereby defining the scale over which preserved land will be distributed. This approach contrasts with many other types of SP studies, which specify the exact location of resource changes and which thereby allow for the estimation of quantifiable distance-decay relationships (Bateman et al. 2006; Hanley et al. 2003).

The issue of policy scale is distinct from concerns related to benefit aggregation, i.e., the number and location of households over which given SP estimates are aggregated (Bateman et al. 2006). Rather, the issue here is the potential transferability of *per household WTP for a specific quantity and type of farmland preservation*, where per household WTP may depend upon on the jurisdictional scale over which a given amount of farmland will be preserved. For example, the per acre, per household WTP for land preservation will likely depend on, among other things, whether a given quantity of land is preserved somewhere within the household's *home community* versus somewhere within the household's *home state*. The distinct issues of benefit aggregation and policy scale, however, both share an association with the spatial dimensions of policy impacts and sampled households.

In sum, estimates of farmland amenity values are often characterized by (1) values linked to preservation at a particular scale and (2) an absence of specifics regarding exact location of parcels in question. These characteristics lead to a variety of possibilities for benefit transfer. For example, a policymaker desiring state-scale welfare estimates for a particular state (e.g., for preservation that will occur within the State of Connecticut) might have access to results estimating WTP for farmland preservation, but only from surveys conducted at the community scale (e.g., for preservation that will occur within specific Connecticut communities, estimated from surveys of community residents). There might also exist state- and community-scale WTP estimates derived from surveys conducted in other states (e.g., Delaware). The resulting out-of-state welfare functions could either be transferred directly, or could be used to derive a mechanism to calibrate for differences in welfare estimates across scale. In the present example, such a mechanism could be used to calibrate the available Connecticut community scale welfare estimate(s) to approximate the desired state scale Connecticut value.

The benefit transfer literature provides little information to assist analysts in determining which of the above possibilities is likely to generate welfare estimates with the least transfer error. Intuition suggests that WTP estimates applicable to different scales could differ greatly, given variations in expected resource proximity and the associated potential for use values. Given this intuition, however, it is unclear whether more valid transfers may be obtained by conducting transfers across different states (rather than across scale within states), or by somehow calibrating welfare estimates across scale within a given state. Such information is critical to the application of benefit transfer to land preservation and other policy contexts in which transfer validity does *not* depend on quantifiable distance-decay relationships (Bateman et al. 2006), because the *ex ante* policy context is not characterized by spatial certainty regarding the exact location of resource changes. Rather, the relevant issue is the size of the political or other jurisdiction (or scale) over which a given amount of resource change will occur.

This issue may also be viewed within the framework of site similarity (Johnston 2007). The benefit transfer literature often addresses similarity in terms of population and site attributes, where site attributes traditionally reflect such factors as the availability of substitutes and complements within a given political jurisdiction (e.g., Barton 2002; Loomis 1992; VandenBerg et al. 2001; Piper and Martin 2001; Rosenberger and Loomis 2001). Policy contexts such as farmland preservation, however, also invoke the notion of similarity in scale. Based on similar reasoning to that found in the distance-decay literature (Bateman et al. 2006), intuition suggests that scale similarity should have a critical role in transfer validity. However, the literature provides no systematic findings to assess the legitimacy of this expectation, or the extent to which scale similarity influences transfer error.

This paper assesses different approaches to benefit transfer of farmland amenity values, with an emphasis on transfers across scale. The analysis emphasizes function-based transfers from choice experiment (CE) results (Morrison and Bergland 2006; Morrison et al. 2002), although parallel issues apply to any transfer of SP welfare estimates. The data are drawn from CE analyses conducted simultaneously in two Northeastern states, addressing farmland preservation within the two states, and at two different jurisdictional scales within each state. Various function-based methods are proposed for benefit transfer, each of which draw from one or more existing studies to provide transferable, function-based estimates of WTP. These include the transfer of like scale per acre WTP values between two different states, denoted a transfer across state. We also consider transfer across scale within a given state. An example would be the use of per acre WTP estimated at the community scale to approximate per acre WTP for a state scale preservation program in the same state. Finally, we develop and test transfer mechanisms that calibrate for differences across scale. Systematic assessment of transfer error provides case-study evidence of transfer validity, and offers insight regarding the most appropriate means to conduct benefit transfer of farmland amenity values.

II. BENEFIT TRANSFER ACROSS SCALE AND STATE—A THEORETICAL FRAMEWORK

To promote concise presentation, the analysis is narrowed to a specific valuation context. The conceptual model, however, may easily be extended to a number of parallel contexts, involving any number of scales or states. In the present case, assume that the researcher has access to SP results from *three* of the following *four* analyses. These include:

1. An empirical valuation function applicable to farmland preservation at jurisdictional *scale x,* derived from studies conducted in one or more communities in *state i*.
2. An empirical valuation function applicable to farmland preservation at jurisdictional *scale y*, derived from a study conducted in *state i*.
3. An empirical valuation function applicable to farmland preservation at jurisdictional *scale x,* derived from studies conducted in one or more communities in *state j*.
4. An empirical valuation function applicable to farmland preservation at jurisdictional *scale y*, derived from a study conducted in *state j*.

Given the assumed absence of one of the above four analyses, analysts must consider the most appropriate ways to analyze the data available from the three existing models to generate the desired, but unavailable, benefit estimate. This might involve transfer across scale within a single state, transfer between states, or some combination of the two.

Formally, assume that the willingness to pay, $WTP_{hk}(\cdot)$, of household h for farmland preservation program k is given by the general function

$$WTP_{hk}(\mathbf{X_k}, S_{hk}, R_{hk}, Y_h), \qquad [1]$$

where

$\mathbf{X_k}$ = vector of variables characterizing outcomes and policy attributes of preservation program k;

S_{hk} = categorical variable identifying the policy scale (e.g., community or state) within which preservation will occur and within which the household resides;

R_{hk} = categorical variable identifying the state in which preservation will occur and within which the household resides;

Y_h = disposable income of household h.

Presume that $R_{hk} = \{ i, j\}$ represents two different states (or other defined regions), e.g., Connecticut ($R_{hk} = i$) and Delaware ($R_{hk} = j$). Further, $S_{hk} = \{x, y\}$ represents two different policy scales within each state within which preserved farmland might be distributed, e.g., community ($S_{hk} = x$) and statewide ($S_{hk} = y$). As an illustrative example of this notation, the WTP of household h, in state i, for farmland preservation at scale x is given by the function $WTP_{hk}(\mathbf{X_k}, R_{hk}=i, S_{hk}=x, Y_h)$. To further condense this notation we adopt a convention in which superscripts are used to identify scale and region, i.e., $WTP_{hk}^{ix} = WTP_{hk}(\mathbf{X_k}, R_{hk}=i, S_{hk}=x, Y_h)$, whereas $WTP_{hk}^{jy} = WTP_{hk}(\mathbf{X_k}, R_{hk}=j, S_{hk}=y, Y_h)$.[3]

For example, assume that the desired but unavailable welfare function is given by WTP_{hk}^{iy}, or WTP for farmland preservation in state i at the scale y. Benefit transfer must somehow capitalize on information in the three available functions $W\hat{T}P_{hk}^{ix}$, $W\hat{T}P_{hk}^{jx}$, and $W\hat{T}P_{hk}^{jy}$, where the hat (^) indicates an empirically estimated function. We consider four possibilities for benefit transfer, based on structural use of information embedded in available preference functions:

1. Function based transfer using only information from $W\hat{T}P_{hk}^{ix}$, or WTP for scale x preservation in state i. This approach conducts a function-based transfer across jurisdictional scale (x to y), within the same state (i). We denote this *transfer across scale*.
2. Function based transfer using only information from $W\hat{T}P_{hk}^{jy}$, or WTP for scale y preservation in state j. This approach conducts a function-based transfer across states (j

to i), but at the same jurisdictional scale (y). We denote this *transfer across state*.

3. Combine information from $W\hat{T}P_{hk}^{jx}$ and $W\hat{T}P_{hk}^{jy}$ to derive a function forecasting the *difference* between WTP at the community (x) and statewide (y) scales for otherwise identical preservation activities, based on information from state j. This function is then used to calibrate $W\hat{T}P_{hk}^{ix}$ in state i to obtain the desired estimate WTP_{hk}^{iy} in the same state. This approach uses results from state j to derive a calibration function predicting the difference between WTP at scale x and scale y. This function is then used to calibrate state i WTP at scale x to scale y. We denote this *cross scale difference calibration*.
4. Combine information from $W\hat{T}P_{hk}^{jx}$ and $W\hat{T}P_{hk}^{jy}$ to derive a function forecasting the *ratio* between WTP at the community (x) and statewide (y) scales for otherwise identical preservation activities, based on information from state j. This function is then used to calibrate $W\hat{T}P_{hk}^{ix}$ in state i to obtain the desired estimate WTP_{hk}^{iy} in the same state. This approach uses results from state j to derive a calibration function predicting the ratio between WTP at scale x and scale y. This function is then used to calibrate state i WTP at scale x to scale y. We denote this *cross scale ratio calibration*.

The four presented methods each incorporate distinct assumptions regarding household welfare from farmland preservation; none has a clear theoretical advantage. In the absence of any theoretical preference, the relative performance of each method becomes an empirical question. We emphasize that parallel transfer methods apply regardless of the specific welfare measure considered unavailable. For example, one could consider WTP_{hk}^{jy} as the desired but unavailable welfare estimate and use the four proposed methods to conduct benefit transfer using existing welfare estimates $W\hat{T}P_{hk}^{jx}$, $W\hat{T}P_{hk}^{ix}$,and $W\hat{T}P_{hk}^{iy}$.

The Random Utility Model

The four above-noted benefit transfer approaches are assessed through their performance in applied function-based benefit transfer, based on CE results.[4] The theoretical model for CEs is derived from the standard random utility specification in which utility is divided into observable and unobservable components (Hanemann 1984). Following [1] above, we assume that the utility of household *h* from preservation program *k* is given by

$$U_{hk}(\mathbf{X_k}, Y_h - Fee_{hk}) = v_{hk}(\mathbf{X_k}, Y_h - Fee_{hk}) + \varepsilon_{hk} \qquad [2]$$

where notation follows (1) above with

Fee_{hk} = cost to the respondent of preservation plan *k*, through a mandatory payment vehicle;

$v_{hk}(\cdot)$ = function representing the empirically measurable component of utility;

ε_{hk} = unobservable component of utility, modeled as econometric error.

As above, we allow utility functions to be conditional on both scale and state. We suppress this notation from the functional specification, and instead incorporate it as superscripts. This provides for four possible utility functions drawn from [2]: $U_{hk}^{jx}(\cdot)$, $U_{hk}^{jy}(\cdot)$, $U_{hk}^{ix}(\cdot)$ and $U_{hk}^{iy}(\cdot)$.

Given the above specification, household *h* chooses among three policy plans, (j=A,B,N). The household may choose option *A*, option *B*, or may reject both options and choose the status quo (neither plan, j=N). A choice of neither plan would result in zero preservation and no preservation policy $\mathbf{X_k}$=0, and zero household cost, $Fee_{hk}=0$. The model assumes that household *h* assesses the utility that would result from available choice options (j=A,B,N) and chooses that which offers the greatest utility. That is, given [2], household *h* will choose plan A if

$$U_{hA}(\mathbf{X_A}, Y_h - Fee_{hA}) \geq U_{hz}(\mathbf{X_z}, Y_h - Fee_{hz}) \quad \text{for } z=B,N, \qquad [3]$$

such that

$$v_{hA}(\mathbf{X_A}, Y_h - Fee_{hA}) + \varepsilon_{hA} \geq v_{hz}(\mathbf{X_z}, Y_h - Fee_{hz}) + \varepsilon_{hz.} \quad [4]$$

If the ε_{hk} are assumed independently and identically drawn from a type I extreme value distribution, the model may be estimated as a conditional logit (CL) model or mixed logit (ML) analog (Maddala 1983; Greene 2003). Estimation of parallel models within scales (x, y) and states (i, j) allows for unique estimates of $\hat{v}_{hk}^{jx}(\cdot)$, $\hat{v}_{hk}^{jy}(\cdot)$, $\hat{v}_{hk}^{ix}(\cdot)$ and $\hat{v}_{hk}^{iy}(\cdot)$, from which welfare estimates may be derived following Hanemann (1984). Benefit transfer assessments draw from welfare measures derived from these estimated functions, either alone or in combination.

III. THE DATA

The data draw on six parallel CE surveys conducted in Connecticut and Delaware. The *Mansfield* and *Preston Land Preservation Surveys* addressed land preservation in these two Connecticut communities, and were implemented over random samples of residents in each community. The *Georgetown* and *Smyrna Land Preservation Surveys* followed a matching approach in two Delaware communities.[5] The *Connecticut* and *Delaware Land Preservation Surveys* represented parallel surveys targeted at statewide preservation in each state, and implemented over random statewide samples.

Survey development required over 18 months of background research, interviews with land use experts and stakeholders, and 14 focus groups (Johnston et al. 1995) including cognitive interviews (Kaplowitz et al. 2004). Extensive pretests were conducted during survey design to ensure that the survey language and format could be easily understood by respondents, that respondents shared interpretations of survey terminology and scenarios, and that survey scenarios captured land use and policy attributes viewed as relevant and realistic by respondents. Focus groups led to a self-administered mail survey, following a CE framework (Adamowicz et al. 1998). Prior to administration of choice questions, the survey provided information on land use and change in respondents' local areas, tradeoffs implicit in land conservation and reminders of the budget constraint. The survey also provided instructions and information for CE questions. This included attribute levels that might occur in choice questions, following guidance in the literature regarding visible choice sets (Bateman et al. 2004).

The CE asked respondents to consider alternative preservation options for hypothetical parcels located in their community or state, depending on the survey version. Respondents were provided with two preservation options that would each preserve farm or forest with varying attributes, "Option A" and "Option B," as well two status quo options. The first status quo option stated, "I would not vote for either program.". The second stated, "I support these programs in general, but my household would/could not pay for either Option A or B." This option was included based on focus group results and prior research (Loomis et al. 1999; Brown et al. 1996) as an outlet for those who might wish to express symbolic support for land preservation, yet nonetheless would not pay for either of the provided options. For purposes of estimation the two status quo options—both indicating a choice of no preservation—were combined into a single category.[6]

Each respondent was provided with three CE questions and was instructed to consider each question as an independent, non-additive choice. Attributes characterized land use outcomes identified by focus groups, interviews, and background research as significant to choices among land preservation options, including type of land preserved, the number of acres, the provision and type of public access, the likelihood of development of unpreserved parcels, and the cost of preservation to the respondent's household. Choice questions also specified the technique that would be used to preserve the land in question, as well as the agent that would be responsible for implementing the technique (Johnston and Duke 2007). Table 1 describes the attributes that distinguished hypothetical preservation options. The fractionated experimental design was constructed by the University of Delaware STATLAB based on a D-optimality criterion (Kuhfeld and Tobias 2005). Table 2 shows attributes and levels in the design.

Surveys were implemented during fall 2005. Surveys were mailed to 3000 randomly selected residents of the four CT and DE communities (750 surveys per community), and 2000 randomly selected residents of the two states (1000 per state). Implementation followed Dillman's Tailored Design Method (2000). Of the 2763 deliverable community surveys, 1136 were returned, for an average response rate of 41.1%. Of the 1834 deliverable statewide surveys 622 were returned, for an average response rate of 33.9%.

Differences Across statewide and Community Scale Choice Experiments

To avoid protests and the potential for respondent confusion, it is critical that policies described in survey scenarios are "perceived as realistic and feasible" (Bateman et al. 2002, p. 116). Hence, while state- and community-scale CEs maintain a high degree of parallelism, some differences were necessary in order to maintain realism. While this presents limitations in terms of statistical analysis (e.g., state and community data cannot be pooled within a single statistical model), it also allows for a more realistic assessment of benefit transfer potential (i.e., it recognizes the fact that realistic policy contexts may differ across scale).

Differences apply to attributes characterizing the number of acres, public access, and program cost (table 2). Preservation acreages, for example, are larger at the state scale. This reflects the fact that statewide farmland preservation programs generally target a greater number of acres than those implemented at the community scale—a fact recognized by focus group participants.[7] Similarly, program cost levels diverge across the two survey scales in response to pretest responses revealing differences in the range of household WTP, and payments perceived as realistic, as related to the range of other question attributes (table 2).

Finally, the public access attribute diverged across state and community scales. The goal for this attribute was to provide access levels interpreted as high, medium, and low by respondents. Focus groups for the community survey revealed that scenarios were viewed as most realistic and salient if they allowed for different types of access on individual preserved parcels (e.g., hunting, walking/biking). At the state scale, in contrast, it was perceived as unrealistic that the state could mandate access for any specific activity on all preserved acres. Hence, the statewide survey characterized public access as the percentage of preserved acres for which access would be permitted (i.e., 100%, 50%, 0%). Additional details are found in table 2.

Given the experimental design, there are 180 unique combinations of land preservation attributes for which per acre welfare measures may be estimated across statewide and community models.[8] As detailed below, per acre transfer errors are calculated and averaged across all possible combinations of preservation attributes. As is typically the case with benefit transfer, this requires reconciliation of variables across study and policy sites (Smith et al. 2002; Johnston et al. 2005). Here, all variables characterizing preservation type are identical across scale (and hence require no reconciliation), except those associated with public access. To reconcile public access variables, we match high, medium, and low categories across scales; this provides an approximation that allows access levels to be compared across state and community scales. Although one would expect that a perfect match among state and community attributes would provide the ideal context for benefit transfer, the current situation presents a more realistic situation in which variable definitions are similar but not universally identical.

IV. THE EMPIRICAL MODEL

The literature offers no firm guidance concerning the most appropriate econometric functional form for the observable component of respondents' utility; while linear functions forms are most common, alternative forms are also used depending on theoretical and empirical considerations (Johnston et al. 2002). In the present case, all scenario attributes (except the number of preserved acres) characterize outcome or policy features of preserved land (table 2). Hence, the influence of these attributes on utility is expected to depend on the number of acres preserved. Given this expected conditionality and to avoid unrealistic model forecasts associated with linear terms in the utility function[9], all non-acreage preservation attributes (land type, preservation method, public access, development risk) enter the model as multiplicative interactions with the number of acres preserved. The remaining attributes, including preservation acres, program cost and an alternative specific constant (ASC) for "neither plan" enter linearly. Hence, following [2], household utility from policy option *k* is given by

$$v_k(\cdot) = \beta_0^{SR}(Neither) + \beta_1^{SR} Acres_k + \sum_{n=2}^{N} \beta_n^{SR}(Acres_k)(X_{kn}) + \beta_{N+1}^{SR}(Fee_k)\,, \qquad [5]$$

where *Neither* is the ASC for "neither plan," $Acres_k$ is the number of acres preserved by option *k*, X_{kn} are attributes of preserved acres, Fee_k is the unavoidable household cost of the plan, and the betas (β) are parameters to be estimated. The superscripts *S* and *R* reflect the fact that parameters β^{SR} may differ across both jurisdictional scale $S_{hk} = \{x, y\}$ and state $R_{hk} = \{i, j\}$.

The model is estimated using mixed logit (ML) with both the coefficient on the ASC (*Neither*) and program cost (*Fee*) specified as random. A normal distribution is assumed for the coefficient on *Neither*; a lognormal distribution is assumed for the coefficient on *Fee*. Sign-reversal is applied to the cost variable prior to estimation. These conventions follow standard approaches for variables of these types (Hensher and Greene 2003; Hu et al. 2005). Preliminary models were also estimated in which the coefficient on preserved acres (*Acres*) was randomized;

many of these models showed no statistically significant improvement over specifications in which a fixed (non-random) coefficient was specified. Hence, a fixed coefficient is specified for this variable. In addition, to simplify subsequent welfare simulations (see below) and prevent convergence difficulties, coefficients on multiplicative interactions were also specified as fixed.

Four final ML models are estimated, following the specification above. Model one is a model estimated from pooled Delaware community data for Smyrna and Georgetown. Model two is estimated from the statewide Delaware data. Model three is estimated from pooled Connecticut community data for Mansfield and Preston. Model four is estimated from the statewide Connecticut data. Specifications are identical for all models, subject to caveats noted in the previous section concerning the differences in variable definitions between state and community surveys. Log-likelihood tests (Mazzotta and Opaluch 1995) fail to reject the appropriateness of pooling individual community data within each state (p=0.38 in DE; p=0.13 in CT). All ML models are estimated using maximum likelihood with Halton draws applied in the likelihood simulation. The statistical fit of ML models is superior to that of their CL counterparts—at $p<0.01$ in all cases—hence ML results are illustrated below.

The focus on benefits transfer implies a comparison of WTP results across models (convergent validity) rather than detailed individual results for each model. As a basis for initial comparison, however, table 3 presents individual ML results for each of the four models. All are statistically significant at $p<0.01$, with statistically significant coefficients conforming to prior expectations, where expectations exist. Relative magnitudes of parameter estimates are also as expected—with the intuitive implication that per acre welfare effects associated with community scale preservation exceed those associated with state scale preservation, *ceteris paribus*.[10] The significance of parameter estimates also varies in many instances across state and community models, providing additional evidence that preferences differ across scale.

Given that the estimated models involve random coefficients, welfare measures (implicit prices, compensating surplus) are simulated following the approach of Hu et al. (2005), following the general framework of Hensher and Greene (2003). We follow Hu et al. (2005) and Johnston and Duke (2007) and present welfare estimates as the mean over the parameter simulation (1000 draws) of median WTP calculated over the coefficient simulation (1000 draws).[11] Additional methodological details are suppressed here for the sake of conciseness, but may be found in Hu et al. (2005), Johnston and Duke (2007) or Hensher and Greene (2003).

V. ASSESSMENTS OF TRANSFER VALIDITY AND ERROR

A variety of tests relevant to the validity of benefits transfer may be conducted. For example, past assessments have included tests of estimated utility parameters and implicit prices (Jiang et al. 2005; Johnston 2007; Morrison et al. 2002). In the present case, we are interested primarily in the relative performance of different approaches to transferring per acre WTP for farmland preservation—a compensating surplus (CS) measure and, as such, the measure most directly relevant to policy (Morrison et al. 2002). Nonetheless, as an initial comparison of model results, table 4 presents implicit prices associated with the four estimated models. We also present two-tailed *p*-values for the null hypothesis of equal implicit prices across scale in Connecticut and Delaware, based on the method of convolutions (Poe et al. 2005).

As shown by table 4, point estimate magnitudes of implicit prices vary to a substantial degree across scale, with community scale implicit prices always larger in absolute value than analogous state scale values. Despite these large point estimate differences, the sometimes large variances of ML parameter estimates (table 3) lead to a failure to reject the null hypothesis of implicit price equality for the majority of implicit prices. However, for 8 out of 24 implicit prices (33%), we reject equality at p=0.10 or better, suggesting that differences in scale, even in the same state, can lead to statistically significant differences in welfare estimates.

Notwithstanding the insight that may be available from implicit prices alone, assessment of CS generally provides a more policy relevant perspective on transfer performance (Morrison et al. 2002). Given the 180 possible preservation types noted above, we consider cases in which each of the four state-scale combinations is treated as the policy site for purposes of benefit transfer. That is, assessments of transfer error are conducted for cases in which each of the four welfare measures (WTP_{hk}^{iy}, WTP_{hk}^{ix}, WTP_{hk}^{jx}, and WTP_{hk}^{jy}) is considered the unknown ,but desired, estimate over all preservation types.

As described above, four empirical, function-based approaches to benefit transfer are tested. These include (a) transfer across scale within the same state, (b) transfer across states at the same scale, (c) cross scale WTP difference calibration using a state-community calibration function estimated in one state, then applied to the second, and (d) cross scale WTP ratio calibration using a state-community calibration function estimated in one state, then applied to the second. The first two methods capitalize on data available from a single source to conduct benefit transfer. The third and fourth, in contrast, use information from the three available estimates to, at least in a sense, triangulate the missing welfare measure.

Given [5] and the associated CS derivation (Boxall et al. 1996), one may easily calculate the per acre WTP difference and ratio within each state R, across scales x and y, as a parametric function of preservation attributes X_{kn}. Drawing from [5], the WTP difference calibration function across state and community scales is thus specified

$$WTP_{dif,k} = \left[{(\hat{\beta}_1^{xR} + \sum_{n=2}^{N} \hat{\beta}_n^{xR}(X_{kn}))} \Big/ {\hat{\beta}_{N+1}^{xR}} \right] - \left[{(\hat{\beta}_1^{yR} + \sum_{n=2}^{N} \hat{\beta}_n^{yR}(X_{kn}))} \Big/ {\hat{\beta}_{N+1}^{yR}} \right]$$

$$= (\hat{\beta}_1^{xR} / \hat{\beta}_{N+1}^{xR} - \hat{\beta}_1^{yR} / \hat{\beta}_{N+1}^{yR}) + ((\sum_{n=2}^{N} \hat{\beta}_n^{xR} / \hat{\beta}_{N+1}^{xR}) - (\sum_{n=2}^{N} \hat{\beta}_n^{yR} / \hat{\beta}_{N+1}^{yR}))X_{kn} \qquad [6]$$

where superscripts x and y denote the two different scales in question. Note that [6] is simply the sum of implicit price differences across scale, where $Acres_k = 1$ (so that WTP reflects marginal value per acre) and X_{kn} are individual attributes upon which WTP per acre is conditional.[12] The function forecasts a unique WTP difference for each preservation type, as characterized by X_{kn}. For the state whose results are used to calculate [6], $WTP_{dif,k}$ perfectly forecasts the WTP difference between the state and community scales, for each possible preservation type.

To conduct benefit transfer using [6], one estimates $WTP_{dif,k}$ for all 180 preservation types defined by the X_{kn}, based on parameter estimates for community (scale x) and state (scale y) models in the first state (considered the study site). Then, to calibrate across scale in the policy site, one either adds or subtracts $WTP_{dif,k}$ from the corresponding estimates of state (scale y) or community (scale x) WTP per acre, respectively, in the second state. The result is a calibration across scale in the policy site, based on a WTP difference function estimated at the study site.[13] This calibration in per acre WTP is unique for each possible preservation type, as characterized by X_{kn}, and presumes that the calibration function [6] is transferable from region $R=i$ to $R=j$.

The fourth method is conceptually analogous to that based on [6], but calibrates according to a *WTP ratio function* across scales x and y. The function is given by

$$WTP_{ratio,k} = \frac{[(\hat{\beta}_1^{xR} + \sum_{n=2}^{N} \hat{\beta}_n^{xR}(X_{kn})) / \hat{\beta}_{N+1}^{xR}]}{[(\hat{\beta}_1^{yR} + \sum_{n=2}^{N} \hat{\beta}_n^{yR}(X_{kn})) / \hat{\beta}_{N+1}^{yR}]} \qquad [7]$$

To conduct benefit transfer using [7], one estimates $WTP_{ratio,k}$ for all preservation types, based on parameter estimates for community (scale x) and state (scale y) models in the first state (considered the study site). To calibrate across scale in the policy site, one then either multiplies or divides $WTP_{ratio,k}$ by the corresponding estimates of state (scale y) or community (scale x) WTP per acre, respectively, in the second state. The result is a calibration across scale in the policy site, based on a WTP ratio function estimated at the study site.

Empirical Results: Implications of Transfer Approach for Transfer Error

Following common convention, transfer error is quantified as a percentage divergence of transfer estimates from an estimated "true" value—as estimated by the CE for the state/scale combination assumed to be the study site in each case (Rosenberger and Stanley 2006). Percentage errors in per acre WTP are presented as an average over all 180 preservation types. Results are illustrated by table 5, along with the true average WTP across preservation types.

As shown by table 5, the choice of function-based transfer approach has crucial implications for transfer error. As an average, transfer across states at the same scale (method two) outperforms other approaches, with an average transfer error of less than 16% in absolute value. In contrast, transfer across scale within the same state (method one) generates an average transfer error in excess of 2000% in absolute value. Although the relative performance of simple across-scale transfer varies depending on whether the community or state scale is the transfer target, there is greater error potential when one transfers across scale. Interestingly, the two simpler methods (across state and across scale transfer) generate the best and worst average performance of the four approaches.

Of the two more complex transfer methods (WTP difference and ratio calibration across scale), ratio calibration outperforms difference calibration, on average. The poorer average performance of difference calibration is related primarily to large transfer errors associated with the prediction of state scale values, with absolute value errors exceeding 3500% in Delaware and 4500% in Connecticut. In contrast, ratio calibration performs more acceptably in these cases—with average transfer errors less than 125% in absolute value.

Standard deviations of transfer error across preservation types suggest similar conclusions (table 4). Standard deviations—and thus the variability in transfer error—are greater, on average, for the three methods involving transfers across scale. Particularly large are standard deviations associated with simple across scale transfer (method one) and difference calibration (method three), particularly when associated with transfer to the state scale. For example, the average standard deviation associated with simple across scale transfer (5941.43) exceeds that associated with across state transfer (399.05) by nearly a factor of fifteen. Even for across state transfer, however, the standard deviation is relatively large. This suggests the potential for substantial transfer errors in individual cases (e.g., types of land or preservation), notwithstanding average transfer errors that may be relatively small. As a result, the superior average performance of across state transfer does not necessarily imply that small errors are to be expected for all types of preservation.

As noted above, results also suggest that transfer errors, at least in percentage terms, depend on whether the policy site (or transfer target) is at the state or community scale. Transfer to community scale preservation generally results in smaller transfer errors than transfer to state scale preservation. These results, however, must be taken in context of the larger WTP baseline from which community scale transfer errors are assessed (table 5). For example, in Delaware, average community scale WTP exceeds analogous state scale WTP by a factor of 27.72; the analogous factor in Connecticut is 103.39. As a result, a given magnitude error will be greater in percentage terms when compared to the smaller welfare measures estimated at the state scale.

Consolidating these results into general findings, a few principal messages emerge. First, results indicate that substantial errors may result if one conducts benefit transfer across scale. That is, scale similarity is a critical determinant of transfer validity. In the present study, across scale transfers generate the worst average performance, by a large margin, with errors commonly exceeding one thousand percent. Even the more complex methods that calibrate across scale fare more poorly than simple transfers between states at the same jurisdictional scale. Based on these findings, it is better to transfer at the same jurisdictional scale than at different jurisdictional scales. Such results suggest caution in the application of welfare results to policies that extend beyond the original valuation scale. For example, based on these results, one should recognize the substantial bias that may occur if one seeks to use land preservation values estimated at the community scale to approximate benefits for a state scale preservation policy.

In the present case, it is unclear whether the poor performance of across-scale transfers is related to differences in proximity between state and community scale preservation, or the fact that statewide preservation programs typically target a greater number of acres—a pattern also reflected in our experimental design (table 2). Both of these factors will encourage larger per acre WTP at the community scale, *ceteris paribus*, and these effects cannot be disentangled effectively using the present data. Hence, it is left for future research to determine the relative contribution of each factor to the observed differences in welfare measures across scale.

The dominance of simple across region transfers over across scale transfers also validates prior findings that geographical proximity alone is insufficient to guarantee transfer validity, when aspects of the policy context differ. Here, transfers within the same state can nonetheless entail substantial point-estimate transfer errors related to differences in policy context (i.e., changes in the scale over which preservation occurs). Similar patterns have been established by

Johnston (2007) in the CE benefit transfer literature and by Piper and Martin (2001) and others in non-CE benefit transfer research—that proximity alone is a poor predictor of transfer error.

Finally, the superior performance of the most common, simple function-based transfer method in the present case provides an interesting—though far from definitive—counterpoint to recent works such as Smith et al. (2002), which propose complex models to improve transfer performance. Such methods invoke a variety of assumptions and mathematical derivations as a means to better ground transfer estimates in underlying welfare theory, thereby rendering transfer more defensible. The current results, in contrast, reflect *lex parsimoniae*. The common across state transfer, invoking perhaps the fewest assumptions, generates the smallest point-estimate transfer error. This is an encouraging, if preliminary finding for analysts who seek to apply benefit transfer, yet lack the ability to conduct more complex analyses. Assessment of the general applicability of such findings, however, including potential applicability to the utility-theoretic transfer methods of Smith et al. (2002) and others, requires additional targeted research and more specific evidence than is available from the present analysis.

VI. CONCLUSION

In many policy contexts, researchers have a variety of options for benefit transfer. These may include transfers across different states or across different jurisdictional scales within the same state. The current literature provides little information to assist analysts in determining which of these possibilities is likely to generate welfare estimates with the least transfer error. This paper seeks to provide insight regarding the validity of across state and across scale benefit transfers, with specific applicability to land preservation values. As in most empirical assessments, there are a variety of analyses that are omitted for the sake of conciseness, and numerous topics that remain unexplored. For example, the present analysis does not seek to adjust welfare estimates for differences in populations across policy scales. Moreover, the analysis applies only to a single state-community scale dichotomy, and to the single policy context of land preservation.

These and other limitations notwithstanding, the analysis offers strong evidence that across state transfers of land preservation values outperform transfers across jurisdictional scale—even when systematic mechanisms are used to calibrate welfare estimates across scale. Put another way, similarity in jurisdictional scale is a critical aspect of benefit transfer. We also find that the simplest, most common function-based method for benefit transfer—the function-based transfer of per acre WTP across states—outperforms other methods on average. It is hoped that such findings, combined with future work, may assist analysts in identifying the most appropriate mechanisms for benefit transfer when original studies are infeasible.

Finally, from the perspective of applied welfare analysis, results suggest that across state benefit transfer can, at least in some instances, provide results with a fairly acceptable degree of transfer error. Simple transfers of CE compensating surplus functions generated relatively modest average errors, at least compared to those found in other resource contexts (Rosenberger and Stanley 2006). This result must be considered, however, within the context of poor performance achieved with across scale transfers. Hence, while function-based transfer of farmland preservation values can be conducted with reasonable error, the specific method used is critical to transfer validity.

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Endnotes

[1] Here, we adopt the more general definition of "stated preference" methods, to include all generally-accepted, direct methods of survey-based valuation (e.g., contingent valuation, choice experiments, etc.).

[2] Loomis (2000) finds that the use of political jurisdictions, rather than the more relevant economic jurisdictions, may lead to significant underestimation of willingness to pay.

[3] To simplify the model, we assume that regardless of preservation scale households only consider preservation that occurs in their home state for state scale welfare assessment, or home community for community scale assessment. While the model could be easily extended to address this possibility, the resulting notation and complexity would detract from the main issues addressed here.

[4] As noted by Johnston (2007), "the ability of choice experiments to explicitly adjust for differences in the attributes of environmental goods or policies provides an increased capacity to adjust for differences between study and policy sites—thereby improving the potential accuracy of benefits transfer (Morrison et al. 2002; Jiang et al. 2005)."

[5] Surveyed communities were selected based on a number of factors, including the presence of similar development pressures, the lack of a major urban center in close proximity, and the existence of large areas of undeveloped land (Johnston and Duke 2007).

[6] This treatment of responses, while simplifying the data, has no substantive impact on model results.

[7] When presented with preliminary surveys showing statewide preservation programs that targeted small numbers of acres, focus group respondents often considered such programs either trivial or unrealistic.

[8] This number is derived by including all combinations of the 4 land types, 5 preservation methods, 3 access types, and 3 risk levels in the design, leading to $4 \times 5 \times 3 \times 3 = 180$ options for which per acre welfare measures are estimated.

[9] For example, a linear specification would predict a fixed utility impact of land attributes regardless of the number of acres preserved. As a result, a linear model would forecast a utility change associated with various land attributes even if zero acres were preserved—a clearly unrealistic outcome.

[10] One might expect lower per acre WTP measures at the state scale both because of the lesser degree of expected proximity to preserved land, and also due to diminishing marginal utility; the statewide survey incorporated much larger acreages, such that the marginal utility per acre would be expected to decline relative to the community scale analysis.

[11] This approach avoids unrealistic mean WTP estimates related to the lognormal distribution of the program cost coefficient, resulting from the long right-hand tail of the distribution (Hensher and Greene 2003). As noted by Hu et al. (2005), there is no strong theoretical preference for either mean or median welfare measures.

[12] As we are calculating WTP for a marginal acre of preservation, and not a new preservation program, we drop the coefficient associated with the alternative specific constant from CS calculations.

[13] For example, assume that the desired but unavailable welfare estimate is per acre WTP at the state scale in Connecticut. One would calculate (6) from available state and community results in Delaware, then use the resulting function to calibrate community scale results available from Connecticut—to obtain a transfer approximation of the desired Connecticut state scale estimate.

Table 1. Variables and Descriptive Statistics

Variable	Description	Mean Value [a] (std. dev)			
		Connecticut Community	Connecticut State	Delaware Community	Delaware State
Neither	Alternative specific constant (dummy) identifying the status quo option.	0.33 (0.47)	0.33 (0.47)	0.33 (0.47)	0.33 (0.47)
Acres	Number of acres preserved.	62.68 (70.01)	4001.18 (3958.15)	63.10 (70.78)	4007.79 (3956.68)
*Acres*Nursery*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the parcel is an active nursery (omitted default is a food or dairy farm).	12.71 (40.54)	840.78 (2441.58)	11.98 (39.52)	846.88 (2464.72)
*Acres*Forest*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the parcel is forest (omitted default is a food or dairy farm).	12.49 (40.21)	825.49 (2433.69)	12.80 (41.22)	777.10 (2347.00)
*Acres*Idle*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the parcel is idle farmland (omitted default is a food or dairy farm).	12.93 (40.73)	798.82 (2364.06)	13.35 (41.37)	793.02 (2355.59)
*Acres*Trust Easement*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that preservation is accomplished through conservation easements, implemented by land trusts, using block grant funds from the state (omitted default is preservation by conservation zoning).	6.64 (29.53)	419.22 (1795.87)	6.73 (30.43)	468.16 (1889.80)
*Acres*State Purchase*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that preservation is accomplished through fee simple purchase of the parcel, implemented by the state (omitted default is preservation by conservation zoning).	20.92 (50.44)	1278.43 (2908.46)	20.85 (50.44)	1291.67 (2914.34)
*Acres*Trust Purchase*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that preservation is accomplished through fee simple purchase of the parcel, implemented by the land trusts, using block grant funds from the state (omitted default is preservation conservation zoning).	20.42 (49.08)	1427.84 (3046.66)	21.31 (50.77)	1326.22 (2951.25)
*Acres*State Easement*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that preservation is accomplished through conservation easements, implemented by the state (omitted default is preservation by conservation zoning).	7.27 (31.59)	412.16 (1751.07)	6.95 (30.53)	450.20 (1847.70)
Acres Moderate*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating	14.38 (42.54)	812.16 (2393.19)	14.56 (43.00)	832.32 (2430.74)

Access	that the preserved parcel would offer moderate levels of public access. This is defined as access for walking and biking in the community survey, and access on 50% of preserved parcels in the state survey (omitted default is no public access).				
*Acres*High Access*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the preserved parcel would offer high levels of public access. This is defined as access for hunting in the community survey, and access on 100% of preserved parcels in the state survey (omitted default is no public access).	12.37 (39.12)	903.14 (2510.63)	12.91 (40.57)	916.67 (2529.63)
*Acres*No Development 30 Years*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the land, if not preserved, would likely remain undeveloped for at least 30 years (omitted default is development likely in less than 10 years).	20.54 (49.27)	1390.98 (3045.31)	22.00 (52.40)	1357.05 (2991.33)
*Acres*Development 10 - 30 Years*	Multiplicative interaction between *Acres* and a binary (dummy) variable indicating that the land, if not preserved, would likely be developed in 10 to 30 years (omitted default is development likely in less than 10 years).	19.81 (48.49)	1225.88 (2817.58)	19.79 (47.95)	1290.99 (2911.03)
Fee	Unavoidable household cost of preservation (state/town taxes and fees), with sign reversal.	-44.27 (63.10)	-75.63 (100.52)	-43.53 (61.79)	-77.49 (102.24)

[a] Includes zeros for the 'neither' option.

Table 2. Attributes and Levels for Choice Experiment Design

Attribute	Levels	
Acres (4 levels)	Community (one parcel) 1. 20 2. 60 3. 100 4. 200	State (multiple parcels) 1. 1,000 2. 5,000 3. 8,000 4. 10,000
Land type (5 levels)	1. Active Farmland a. Nursery b. Food Crop c. Dairy or Livestock 2. Farmland (currently idle) 3. Forest	
Policy technique and implementing agency (5 levels)	1. Preservation Contracts a. By State b. By Land Trusts using Block Grants 2. Outright Purchase a. By State b. By Land Trusts using Block Grants 3. Conservation Zoning	
Public access (3 levels)	Community (one parcel) 1. No Access Allowed 2. Access for Walking & Biking 3. Access for Hunting	State (multiple parcels) 1. No Access Allowed 2. Access on 50% of Parcels 3. Access on 100% of Parcels
Development risk (3 levels)	1. Development likely in less than 10 years if not preserved 2. Development likely in 10-30 years if not preserved 3. Development NOT likely within the next 30 years	
Cost (6 levels)	Community 1. $5 2. $15 3. $30 4. $50 5. $100 6. $200	State 1. $10 2. $30 3. $50 4. $100 5. $200 6. $300

Table 3. Mixed Logit Results

	Delaware Community	Delaware State	Connecticut Community	Connecticut State
Neither (ASC)	-0.93298	-0.720424	-0.28075	-2.327030
	(0.20365)***	(0.293863)***	(0.16271)*	(0.422373)***
Fee (lognormal, sign reverse)	-3.72041	-4.520530	-4.88730	-4.474820
	(0.24206)***	(0.323230)***	(0.17596)***	(0.249562)***
Acres	-0.00237	-0.000009	0.00016	-0.000068
	(0.00207)	(0.000043)	(0.00145)	(0.000042)
*Acres*Nursery*	-0.00106	-0.000027	-0.00285	-0.000043
	(0.00123)	(0.000026)	(0.00098)***	(0.000029)
*Acres*Forest*	0.00008	-0.000006	0.00064	0.000038
	(0.00124)	(0.000029)	(0.00104)	(0.000029)
*Acres*Idle*	0.00066	-0.000011	-0.00104	0.000006
	(0.00126)	(0.000027)	(0.00103)	(0.000029)
*Acres*Trust Easement*	0.00171	0.000098	0.00223	0.000168
	(0.00226)	(0.000047)**	(0.00171)	(0.000047)***
*Acres*State Purchase*	0.00421	0.000089	0.00219	0.000053
	(0.00189)**	(0.000041)**	(0.00156)	(0.000045)
*Acres*Trust Purchase*	0.00096	0.000091	0.00334	0.000085
	(0.00197)	(0.000042)**	(0.00167)**	(0.000044)**
	0.00573	0.000091	0.00284	0.000096
*Acres*State Easement*	(0.00209)***	(0.000050)*	(0.00174)*	(0.000053)*
	0.00803	0.000086	0.00773	0.000120
*Acres*Moderate Access*	(0.00156)***	(0.000030)***	(0.00126)***	(0.000038)***
	0.00609	0.000072	0.00155	0.000126
*Acres*High Access*	(0.00151)***	(0.000029)***	(0.00122)	(0.000034)***
*Acres*No Development 30*	-0.00061	-0.000106	-0.00192	-0.000075
Years	(0.00097)	(0.000023)***	(0.00084)**	(0.000023)***
*Acres*Development 10 - 30*	-0.00149	-0.000019	0.00039	-0.000034
Years	(0.00116)	(0.000022)	(0.00085)	(0.000025)
	1.53389	1.784680	1.98357	2.788900
std NE	(0.39456)***	(0.493429)***	(0.19766)***	(0.525565)***
	2.56899	2.677350	1.73848	2.569740
std Cost	(0.30388)***	(0.443947)***	(0.29246)***	(0.285337)***
Log-Likelihood Chi-Square	630.01***	444.83***	557.26***	400.82***
Pseudo-R^2	0.20	0.21	0.14	0.22
N	4308	2952	5625	2550

Note: Single (*), double (**) and triple (***) asterisks denote p-values of 0.10, 0.05 and 0.01, respectively.

Table 4. Implicit Prices and Policy Scale: Results for Delaware and Connecticut

Implicit Price[a]	**Farmland preserved somewhere within—**			**Farmland preserved somewhere within—**		
	Connecticut Community	Connecticut State	*p*-value[b]	Delaware Community	Delaware State	*p*-value[b]
Acres	0.0246	-0.0065	0.88	-0.1014	-0.0013	0.24
*Acres*Nursery*	-0.3783	-0.0038	0.00	-0.0429	-0.0024	0.44
*Acres*Forest*	0.0810	0.0035	0.58	0.0018	-0.0006	0.94
*Acres*Idle*	-0.1458	0.0006	0.30	0.0245	-0.0009	0.60
*Acres*Trust Easement*	0.2958	0.0157	0.22	0.0733	0.0097	0.52
*Acres*State Purchase*	0.2817	0.0052	0.16	0.1747	0.0090	0.04
*Acres*Trust Purchase*	0.4385	0.0081	0.06	0.0396	0.0092	0.70
*Acres*State Easement*	0.3882	0.0089	0.10	0.2462	0.0091	0.00
*Acres*Moderate Access*	1.0500	0.0108	0.00	0.3412	0.0083	0.00
*Acres*High Access*	0.2191	0.0115	0.20	0.2595	0.0068	0.00
*Acres*No Development 30 Years*	-0.2598	-0.0068	0.02	-0.0254	-0.0102	0.74
*Acres*Development 10 - 30 Years*	0.0535	-0.0031	0.62	-0.0628	-0.0018	0.20

[a] Implicit prices are calculated as the mean over the parameter simulation (1000 draws) of median implicit prices calculated over the coefficient simulation (1000 draws), following Hu, Veeman and Adamowicz (2005) and Johnston and Duke (2007).

[b] Illustrated two-tailed p-values are for the null hypothesis of equal implicit prices across scale (i.e., state versus community), and are derived using a complete combinatorial convolutions approach (Poe et al. 2005).

Table 5. Transfer Errors Across Scale and/or Region: Comparison of Four Methods

		Average Transfer Error[a]			
Assumed Policy Site for Benefit Transfer	Average per Acre WTP – actual [b]	Method One: Across Scale Within State	Method Two: Across States at Same Scale	Method Three: WTP Difference Calibration	Method Four: WTP Ratio Calibration
Connecticut Community (WTP_{hk}^{ix})	\$0.549 (\$0.528)	-99.56% (8.54)	-77.15% (500.49)	-77.10% (503.55)	-106.27% (735.12)
Connecticut Statewide (WTP_{hk}^{iy})	\$0.005 (0.008)	6278.46% (20312.78)	-29.34% (204.97)	4519.03% (17697.65)	-123.99% (1905.24)
Delaware Community (WTP_{hk}^{jx})	\$0.172 (0.176)	-98.93% (18.83)	48.57% (787.50)	50.93% (782.49)	-123.99% (1905.24)
Delaware Statewide (WTP_{hk}^{jy})	\$0.006 (0.007)	2341.05% (3425.57)	-5.38% (103.25)	-3688.17% (7147.95)	-106.27% (735.12)
Average Error	--	2105.25%	-15.83%	201.18%	-115.13%
Average Standard Deviation	--	5941.43	399.05	6532.91	1320.20

[a] Transfer error in per acre mean of median WTP averaged over N=180 possible preservation types. Values in parentheses are standard deviations over preservation types. In two instances, a single outlier observation was omitted due to an actual WTP estimate that approximated zero, leading to nearly infinite percentage errors. For these two cases, N=179.

[b] Calculated as a mean (and standard deviation) in estimated mean of median WTP across all preservation types.