Private Options to Use Public Goods
Exploiting Revealed Preferences to Estimate Environmental Benefits

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Summary

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PRIVATE OPTIONS TO USE PUBLIC GOODS
EXPLOITING REVEALED PREFERENCES TO ESTIMATE ENVIRONMENTAL BENEFITS

Lori D. Snyder, Robert N. Stavins and Alexander F. Wagner*

We develop and apply a new method for estimating the economic benefits of an environmental amenity. The method fits within the household production framework (Becker 1965), and is based upon the notion of estimating the derived demand for a privately traded option to utilize a freely-available public good. In particular, the demand for state fishing licenses is used to infer the benefits of recreational fishing. Using panel data on state fishing license sales and prices for the continental United States over a fifteen-year period, combined with data on substitute prices and demographic variables, a license demand function is estimated with instrumental variable procedures to allow for the potential endogeneity of administered prices. The econometric results lead to estimates of the benefits of a fishing license, and subsequently to the expected benefits of a recreational fishing day. In contrast with previous studies, which have utilized travel cost or hypothetical market methods, our approach provides estimates that are directly comparable across geographic areas. Further, our results suggest that the benefits of recreational fishing days are generally less than previously estimated.

1. INTRODUCTION

When considering a number of disparate environmental and natural resource issues, policy makers may wish to have estimates of the economic value of a day of recreational fishing. Indeed, in the past, such estimates have been used in analyses of the consequences of: new dams and reservoirs, improvements in water quality, cleanups of abandoned hazardous waste sites, and reductions in the magnitude of global climate change. Virtually all of these estimates have been made with one of two methods: contingent valuation, a direct survey approach employing hypothetical constructed markets; or travel-cost, an indirect market-based method. The use — more broadly — of the first of these approaches has generated considerable controversy within economics;[1] and both approaches require large quantities of geographically specific data.

In this context, it is of interest to have an independent set of estimates — based upon a conceptually distinct, revealed-preference approach — of the economic benefits of a recreational fishing-day. Our methodology differs from previous studies in two important respects. First, we estimate willingness-to-pay for a recreational fishing day from observed behavior regarding the purchase of fishing licenses, rather than observed behavior regarding travel to sites or stated

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[1]For an overview of the controversy, see: Portney 1994; Hanemann 1994; and Diamond and Hausman 1994.

preferences regarding those sites. Second, the two existing approaches[2] use detailed micro-data (of observations of opportunity costs of travel or respondents' explicit estimates of willingness-to-pay) to develop benefit estimates specific to particular bodies of water. This is both their advantage and disadvantage. In contrast, the approach developed in this paper uses aggregate data at the state level to derive estimates — in a national modeling framework — of state averages of recreational benefits.[3] As a result, our state estimates are directly comparable among one another, allowing inferences to be made about relative recreational benefits across geographic areas with more confidence than is possible based on previously available methods.

Given these methodological differences, it is perhaps not surprising that our results differ — in some cases significantly — from those previously available. In particular, our estimates are generally lower — in some cases, much lower — than those from previous studies.

1.1 Overview of Method

Throughout the United States, a state fishing license is required for recreational fishing on any and all bodies of water, with the exception of privately own ponds. Thus, apart from the possibility of illegal fishing activity, to which we return below, a license is a necessary condition for deriving benefits from a day of recreational fishing. Likewise, apart from the relatively rare urbanite who may enjoy displaying to others an (unused) fishing license, experiencing some fishing days is a necessary prerequisite for deriving benefits from owning a fishing license.[4]

Building upon the household production approach to consumer behavior (Becker 1965), Bockstael and McConnell (1983) identified the conditions under which empirical knowledge of the demand function for a private, market good could be used to infer the benefits derived from a related public good. In a simple model,[5] let X be the number of fishing days experienced, L a fishing license, and Z a composite of other goods and services. If utility is defined by the function,

$$U = U(X, L, Z) \tag{1}$$

then the above situation can be represented by the following pair of marginal utility relationships:

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[2]Other direct, revealed-preference methods that have been used for examining environmental amenities and that require detailed micro-data — hedonic property and wage models — have not been applied to estimating the value of recreational fishing days (Freeman 2003). In principle, a sport fishing demand function could also be estimated in some cases by drawing on data from pay-for-use facilities, such as private, managed trout ponds, where users are charged for access or use (Vaughan and Russell 1982).

[3]Seneca and Davis (1976), in an analysis of the factors affecting participation in recreational activities, carried out a county-level, cross-sectional econometric analysis of the factors affecting fishing license sales in West Virginia in 1970. Because there was no variation in license prices in the cross section, price could not be included as an explanatory variable.

[4]We return later to the possibility that the license provides its owner with the *option* to go fishing, and thereby that simply *expecting* to go fishing is all that is required for a person to derive benefits from owning a license.

[5]For a detailed theory of the utilization of recreational fisheries, see: Anderson 1993.

$$\frac{\partial U\ (X, 0, Z)}{\partial X} \quad = \quad 0 \tag{2}$$

$$\frac{\partial U\ (0, L, Z)}{\partial L} \quad = \quad 0 \tag{3}$$

defining what McConnell (1992) terms "joint weak complementarity." Thus, we can employ information about peoples' revealed valuation of fishing licenses, measured by the appropriate area under the respective demand functions, to draw inferences regarding their revealed valuation of (expected) recreational fishing days (Smith 1991).

The first step is to estimate econometrically a set of demand functions for state annual recreational fishing licenses. By measuring the appropriate area under the (state-specific) inverse demand function, we can estimate the average benefits per capita of fishing licenses. Further manipulation leads to an estimate of the average benefits of fishing licenses *per license* (again, specific to states and years). From this we derive a revealed-preference estimate of the expected value of a recreational fishing day. This value can be compared with estimates derived in previous studies by contingent valuation, travel cost, or other methods.

1.2 Preview of the Paper

In Part 2, we describe our data, and in Part 3, we describe the econometric analysis, including the results from ordinary least squares (OLS) and instrumental variables (IV) regressions. In Part 4, we use the econometric results from the IV equations to derive estimates of average expected recreational fishing day values, and we compare these estimates with results from previous studies. In Part 5, we review methodological implications and reflect on the numerical results.

2. DATA

Recreational fishing licenses are sold by all fifty states. In all cases, prices are administratively set by state governments, and licenses are sold without limit. This study focuses on a panel of licenses sold in 48 states[6] over a 15-year period (1975 to 1989).[7] We aggregated the numerous types of fishing licenses that exist into ten categories. All states offered both resident and nonresident

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[6]Hawaii is excluded from the analysis because the structure of license demand does not match the rest of the United States, and Tennessee is excluded because it does not sell a fishing license *per se*, but a "Sportsman License," which can be used for both hunting and fishing.

[7]Since our motivation and contribution is largely methodological in nature, we have chosen a time period for our analysis for which a large number of previous studies exist, the results of which can be compared with our own. In any event, it is highly unlikely that fishable waters have changed in the interim to such degree as to change significantly our relative and absolute estimates.

licenses, the former at lower prices. Resident annual licenses were by far the most popular type, with sales of more than 257 million over the 15-year sample period, representing about two-thirds of all licenses sold (Table 1).[8] Second in numerical importance were resident "combination licenses" that allow for both hunting and fishing during a given year.

Various duration short-term fishing licenses were also available to residents in many states; these allowed for as little as a single day or as much as two weeks of fishing, but — in total — made up only 3 percent of all resident fishing license sales. In contrast, nonresident license sales were much more heavily weighted toward short-term permits. In fact, about 65 percent of all nonresident fishing license sales during the sample period were of short-term licenses (Table 1).

There was substantial variation in aggregate and per capita sales of the various types of licenses. The variation was greatest across states, but was also significant within states over time. For example, in the final year of the sample, 1989, resident annual license sales ranged from about one percent to over 25 percent of state population. There was also considerable variation in (real) license prices across states and over time. In the case of resident annual licenses, for example, the range in the sample was from a minimum of \$3.79 to a maximum of \$32.80[9] (Table 2). In the final year of the sample, the range was from \$7.63 (Minnesota) to \$26.73 (Colorado). The variation was even greater for some of the other license categories. For example, in 1989, the minimum price of a nonresident annual license was \$16.66 (South Dakota) and the maximum was \$69.44 (California).

In addition to the license price and quantity information, data were assembled on various demographic, political, and environmental variables that were thought to be relevant to license demand or necessary for establishing the links between license demand and participants' expected benefits of a recreational fishing day. These data are summarized in Table 2 and discussed in the next section.

3. ECONOMETRIC ANALYSIS OF LICENSE DEMAND

Since license prices are set administratively by states, and licenses are sold without limit, the quantity sold at various prices traces out a demand function if the administratively set prices are exogenously determined and if any relevant variables that are omitted are uncorrelated with the license price.[10] The first set of variables — in addition to the license price itself — that would seem to be relevant are the prices of major substitutes. In this analysis we focus on the demand for resident annual fishing licenses, and therefore the relevant substitute prices include the price of

[8]Two states — Montana and Wyoming — required the purchase of a "conservation stamp" in addition to a fishing license. The cost of the stamp was added to the price of a license.

[9]These and all other monetary amounts in this paper are expressed in year 2000 dollars.

[10]Under these conditions, the observed price-quantity combinations are the intersections of infinitely elastic license supply functions (one for each price) and an assumed downward-sloping license demand function. This situation contrasts with hunting licenses, which are *not* sold without limit in most states. At administered prices, there are typically shortages, which are dealt with in various ways, including random lotteries and waiting lists. Because of this difference, the approach described in this paper is not directly applicable to valuing a recreational hunting day.

resident short-term fishing licenses, the price of resident combination fishing and hunting licenses, and the price of nonresident licenses in adjacent states.[11]

Presumably, the characteristics of demanders are also relevant, and we therefore included the following demographic variables in the resident license demand estimation: median family income; mean years of education; and the share of the population living in urban areas. Finally, the nature of available recreational fishing resources in states should affect demand for state fishing licenses. We dealt with this in two ways. First, we included a variable that measures acres of "fishable waters" per state,[12] but this treats all fishing resources as being homogeneous in terms of the experiences they offer. Clearly, this is not correct. An acre of pristine, high-quality Colorado mountain stream is not equivalent to an acre of Ohio reservoir. The omitted variable — quality of fishing waters — is likely to be important and may be correlated with license price, hence causing biased estimates of demand elasticity. The problem presented by such unobserved quality is not insurmountable, because although quality variation is dramatic across states, quality variation within states over time is trivial by comparison. Hence, we can model this unobserved factor as a fixed effect.

This leaves one concern regarding the possibility of inferring a true demand relationship from econometric estimates: is it supply, demand, or some combination of the two that is being observed? In other words, are price and quantity simultaneously determined, or are prices *exogenously* set by states? A reasonable first approximation is that the administered prices are set exogenously by state officials, and so we begin with ordinary least squares (OLS) estimates. But it is not difficult to posit theories of administered prices that support the notion that these prices are endogenous. Hence, we follow the OLS estimates with a set of specifications in which we treat the license price as endogenous, and estimate the relationships with instrumental variable (IV) methods.

3.1 Ordinary Least Squares (OLS) Estimation

Resident annual licenses comprise approximately two-thirds of all fishing licenses sold in the United States.[13] For resident license demand, the dependent variable was expressed as sales per capita.[14] Fixed effects were employed to control for constant differences among states in the quantity and quality of their recreational fishing resources. Thus, for various specifications of the demand for resident annual licenses, we have:

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[11]A problem arises in specifying which state licenses are relevant as substitutes. We constructed a variable that is a weighted average of prices of specified types of nonresident licenses in adjacent states and Canadian provinces.

[12]Note that this variable varies not only across states, but also over time, reflecting both development of new reservoirs and changes in water quality.

[13]To whatever degree the holders of annual licenses have greater or lesser fishing-day valuations than holders of other categories of fishing licenses, the eventual results will tend to over or under-estimate average state valuations.

[14]These equations were also estimated with sales as the dependent variable and state population as an independent variable; the estimated parameters on population were not significantly different from 1.0 (in the primitive equations). An additional reason for estimating the demand equations in per capita terms was to reduce potential problems associated with heteroskedasticity in the error terms.

$$\frac{Q_{it}}{N_{it}} = f\left(P_{it}, P_{it}^{S1}, D_{it}^{S1}, P_{it}^{S2}, D_{it}^{S2}, P_{it}^{S3}, D_{it}^{S3}, P_{it}^{NR}, F_{it}, U_{it}, E_{it}, Y_{it}, D_i, D_t, \varepsilon_{it}, \beta\right) \tag{4}$$

where Q_{it} = quantity of sales of resident annual license in state i in year t;

N_{it} = population of state i in year t;

P_{it} = price of resident annual license in state i in year t;

P_{it}^{S1} = price of short-term, type 1 (1-3 day) resident license in state i in year t;

D_{it}^{S1} = dummy variable which equals unity if a short-term, type 1 resident license is not offered in state i in year t, and otherwise equals zero;

P_{it}^{S2} = price of short-term, type 2 (4-9 day) resident license in state i in year t;

D_{it}^{S2} = dummy variable which equals unity if a short-term, type 2 resident license is not offered in state i in year t, and otherwise equals zero;

P_{it}^{S3} = price of short-term, type 3 (10-15 days) resident license in state i in year t;

D_{it}^{S3} = dummy variable which equals unity if a short-term, type 3 resident license is not offered in state i in year t, and otherwise equals zero;

P_{it}^{NR} = average price of adjacent state nonresident annual licenses for state i in year t;

F_{it} = area of fishable waters (acres) in state i in year t;

U_{it} = share of population living in urban areas in state i in year t;

E_{it} = mean years of education of population in state i in year t;

Y_{it} = median family income in state i in year t;

D_i = state fixed effects;

D_t = annual fixed effects;

ϵ_{it} = an independent, but not necessarily homoscedastic error term;

β = parameters to be estimated.

The results of estimating the fixed effects models of demand for resident annual fishing license are reported in Table 3 for three functional forms: linear; multiplicative (log-log); and semilog. For each functional form, we report two specifications: one includes the prices of all relevant substitutes as explanatory variables; and the other includes only the price of short-term Type 1 licenses plus dummy variables for each year.

In general, estimated own price effects were negative and statistically significant, and substitute price effects were positive, as expected.[15] The parsimonious specification that included

[15]But when the full menu of substitute prices were included, some of the respective parameters were insignificant and negative. The prices of nonresident licenses in neighboring states performed particularly poorly. This could be because

only short-term Type 1 resident licenses as substitutes consistently yielded positive and statistically significant coefficients (for all three functional forms).[16] Goodness-of-fit statistics were reasonably good for these fixed-effects models, with R^2 on the order of .15 to .23; not surprisingly, the complete models — including the fixed effects — explain a much greater share of the observed variance.[17]

It is simplest to consider the results from the multiplicative (log-log) specification, because the estimated coefficients can be interpreted as elasticities. The own-price elasticity was consistently – 0.19 to – 0.20.[18] The cross-price elasticities (of substitute licenses) were all positive in the parsimonious specification, and quite small.[19] The attempt to capture (partially) resource-quality effects with the fishable waters variable met with limited success. The variable was consistently positive and statistically significant, but the elasticities were small. Presumably, much of the variation in the quality of fishing resources across states was picked up by the fixed effects. Finally, the demographic variables seem to have had some effect on fishing license demand, but while the signs were consistent across specifications, the statistical significance and magnitude of the effects varied. In several specifications, income was positive and statistically significant. In other specifications, years of education was negative and significant.

3.2 Potential Problems

These results raise two major concerns: the effect of illegal fishing activity on the estimates; and the possibility of endogeneity of license prices. Since the purpose of econometrically estimating fishing license demand is to derive implied valuations of expected recreational fishing days, it is necessary to consider the implications of illegal fishing, that is, fishing without a license.[20] This, by

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the simple arithmetic average of neighboring state prices does not correctly capture the role that neighboring state fishing opportunities play in the demand for resident annual licenses. If sufficient data were available, it would be preferable to allow the econometrics to determine the appropriate weighting of the neighboring state prices.

[16]The latter specification includes dummy variables for each year in the sample. On average, nominal license prices changed only 2.5 times per state over the 15-year sample period. Therefore, much (but not all) of the intertemporal variation in prices consists of gradually declining *real* prices. If, at the same time, sales were drifting upward, this would yield a negative correlation, but not one due to price-quantity demand effects. The yearly dummy variables were included to examine this potential problem. However, their inclusion did not materially affect the results, and so it appears unlikely that the observed negative price elasticities were due to such spurious correlation.

[17]A Hausman specification test consistently rejected the hypothesis that state-level variation could be adequately modeled as a random effect.

[18]The estimated elasticities at the mean from the other specifications are roughly consistent with these constant elasticities of demand.

[19]Not all states offer all categories of substitute (short-term) licenses during all time periods. In a sense, the "prices" of these non-existent licenses are infinite. The various specifications allow for the effect of some type of license not being available through the inclusion of dummy variables, D_{it}, where $(1-D_{it})$ is interacted with the respective license price, so that for each observation either a substitute price effect or a lack-of-substitute effect is estimated. In theory, both should be positive, which they consistently were.

[20]If we were concerned with the demand for fishing licenses per se, then illegal fishing would not be a problem for the econometrics; indeed, in that case it is important to exclude illegal fishing, as the data implicitly do. Why might one

itself, need not be a problem given the approach that is taken below to derive valuation from license demand, but it can lead to bias in the econometric estimates.

A theoretically desirable way of treating illegal fishing in the license demand equation would be to allow for this particular substitute activity. Hence, we would want to include as an explanatory variable the "price" of illegal fishing, which may be thought of as the magnitude of fines multiplied by the probability of being fined. Unfortunately, these fines are typically set by courts, not by statute or regulation, and data even for proxies of the probability of being fined (enforcement levels) are rarely available.[21] Although it is reasonable to assume that the true "price of illegal fishing" is positively correlated with the demand for fishing licenses, it is much less clear how it is correlated with license prices, if at all. Hence, omitting this variable may not seriously bias the elasticity estimates.[22]

There is also the possibility that license prices, administratively set by governments, are endogenous, that is, that causality runs not only from price to quantity, but vice-versa. How could this be the case? One potential source of such a causal linkage would be state budgets. States might seek to set license prices at levels that cover annual budgets of fish and game services, assumed to be more or less fixed over time. In the time series for a single state, this could yield a negative correlation between quantity of license sales and administered price (recognizing that this assumes that states can predict sales).

Another potential explanation for a spurious, negative price-quantity correlation is associated with persons sorting themselves for residence among states. People with strong preferences for fishing may be expected to exert political pressure to keep license prices low. If people with strong preferences for fishing move to relatively good fishing states (or develop preferences for fishing as a result of having been born and raised in such a state), then states with large quantities of license sales could tend to have relatively low prices, suggesting a source of spurious, negative cross-sectional correlation. We allow for price endogeneity by identifying a set of instruments for license price, and using instrumental variables in what is essentially a reduced form approach.[23]

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be interested in license demand in and of itself? One reason is that such an analysis can provide the relevant elasticities for examining revenue and other effects of fishing license taxes of various forms.

[21]To whatever degree these factors and their effects vary across states but are constant over time, they are picked up by the state fixed effects.

[22]A more serious problem arises, however, if illegal fishing increases when license prices increase, perhaps as a form of protest. If this is the case, then demand responsiveness will be overestimated and the benefits of licenses will be underestimated.

[23]A third potential problem with the fixed effects estimation results presented in the previous section is that the fixed effects model allows the intercept to vary, but other demand parameters are constrained to be equal across states. The data do not permit estimation of separate demand functions for each state.

3.3 Instrumental Variable (IV) Estimation

To address possible price endogeneity, we wanted an instrument or set of instruments that would be correlated with resident annual license prices, but uncorrelated with the error term (uncorrelated with license sales), that is, exogenous to the demand for fishing licenses. Since license prices were administratively set in each state, a reasonable set of instruments would be ones that are indicative of bureaucratic and political proclivities of states, such as the size of government, and the degree of regulation and taxation. In particular, it would be desirable to have instruments that represent states' (possibly changing) proclivities to employ user fees as opposed to taxes and regulations. These proclivities might be correlated with the administrative price of the fishing licenses, but it is less likely that they would be correlated with the quantity of fishing licenses sold.

The set of instruments used in this analysis were: cigarette taxes (cents per package); motor fuels taxes cents per gallon); general sales taxes (percent); and state expenditures (millions of dollars). The IV regression results are reported for the parsimonious specification for all three functional forms in Table 4. The results were robust to different specifications, including changes in the list of substitute prices and changes in the demographic variables. The parameters on the price and fishable waters variables are all of the logical sign and statistically significant, although several are quite small in magnitude. The demographic variable parameters are all of the expected sign, although some are not statistically different from zero. The own-price elasticity of demand is consistently greater (in absolute value) in the IV estimates than in the OLS estimates.[24]

4. ESTIMATING THE VALUE OF A RECREATIONAL FISHING DAY

Three steps were required to derive the (state and year specific) average expected value of a recreational fishing day from the econometrically estimated demand functions for fishing licenses: derive average benefits of fishing licenses per capita from an estimated demand function; calculate average benefits per licensee; and estimate average expected value of a recreational fishing day.

4.1 Estimating Average Benefits of Owning Fishing Licenses from the Demand Function

To derive the average benefits of a fishing license (in per capita terms), we begin with an equation for which the parameters have been estimated econometrically. Consider the instrumental variables setup reported in Table 4. From the estimated demand curve, we can obtain point

[24]This suggests that whatever endogeneity underlies these results, it is not due to either of the linkages posited above, both of which suggested a negative bias for the OLS results. A Hausman specification test can be used to test whether the coefficients in the OLS and IV regressions are statistically different. Using this test, we cannot reject the null hypothesis that the coefficients are the same, that is, we cannot reject the hypothesis that OLS is the correct model. But the Hausman test looks at the weighted difference of all the coefficients, and we are only concerned with whether the price coefficient is statistically different. A different test regresses the license price on the instrumental variables and then includes the predicted price from this first-stage regression as an explanatory variable (in addition to the licence price itself) in the quantity regression. If the coefficient on the predicted price variable is statistically significant, then it is drawing variance from the error term that would otherwise be attributed to the price variable. This would be what we would expect if the price variable was endogenous. We found that the predicted price variable was consistently insignificant when included in the quantity regression. Therefore, we cannot reject the hypothesis that the price variables is exogenous.

estimates as well as uncertainty estimates (in the form of a variance-covariance matrix). We begin with calculations based on the point estimates.

We first set all variables — with the exception of the (annual resident) license price and the dependent variable (sales per capita) — equal to their actual values for a given state and year. Also, we set all parameters at their econometrically estimated values. The inverted form of the resulting equation (that is, the inverse demand function) is then integrated between the actual per capita demand (sales per capita) and zero (or an appropriate cutoff value for the log-log specification), yielding an estimate of the (revealed) benefits per capita of fishing licenses for each state and year. Multiplying by the state's population produces an estimate of the total benefits of licenses; and dividing this by license sales yields an estimate of the average revealed benefits of owning a fishing license *per licensee*:

$$B_L \quad = \quad \left[\int_c^{q_{it}} f(\hat{\alpha}_{it}, \hat{\beta}_0, q)\, dq \right] \cdot \frac{N}{Q} \tag{5}$$

where $f(\cdot)$ = inverse demand function;

q_{it} = per capita sales of resident annual licenses in state i in year t;

c = appropriate cutoff (zero for the linear and semi-log specifications of the demand curve);

α_{it} = the fitted value from setting all variables — other than (annual resident) license price and the dependent variable — equal to their actual values for state i in year t, and all parameters, including the relevant fixed, state effect, at their econometrically estimated values;

β_0 = the estimated own-price elasticity of demand; and

B_L = average benefits of owning a fishing license (per licensee).

We primarily employ the log-log specification of the demand function for benefit-estimation purposes, because we believe on theoretical grounds that the demand for fishing licenses is unlikely to be linear, and because the log-log form consistently provides a better fit than the semi-log functional form. With this specification, however, it is impossible to integrate the demand curve between zero and actual per capita sales, q_{it}, and so we need to identify a limit for purposes of integration. We examined a variety of such cutoff values, including \$33,[25] \$100, \$200, and \$500. The estimates implied by the log-log specification with a cutoff of \$200 are similar to those implied by the semi-log specification, and so we later focus on these results. It is important to note, however, that the *relative* magnitudes of the estimated benefits of recreational fishing across states do *not* change with different cutoff values.

[25]This cutoff corresponds to the highest price actually charged for a license in our sample (\$32.80), and provides a very conservative estimate of benefits.

4.2 Estimating the Expected Value of a Recreational Fishing Day

Since an annual fishing license is essentially an option to purchase (through direct expenditures plus opportunity costs of time) some number of days of recreational fishing experience (up to the total number of days in the season), it would seem that the stochastic relationship between a fishing license and the experiences it can facilitate would bring forth an important component of "option value." After all, no one knows with certainty how many days they will be able to go fishing during a season. Two conditions, however, essentially undue the option value, and make it possible to infer user value directly. First, annual license sales continue throughout the season, up until the very last day. Hence, there is no necessity to purchase a license *before* the time of one's first expedition. Second, numerous short-term licenses (substitutes) are always available.[26]

Thus, if short-term licenses did not exist *and* if there were an early deadline for purchasing annual licenses, then it could be argued that any valuation linked with licenses would include both user and option value. In the absence of those conditions, however, it seems more reasonable to assume that the benefits of a fishing license are linked with expected use value. If we assume that consumers are risk neutral in regard to their fishing license purchases and that the duration of a fishing season is short enough that discounting is not a significant issue, then we can derive the approximate conditional value of an expected fishing day in a very direct manner from the license demand evidence. First, we can say that:

$$B_L \quad = \quad E\left[B_{FD} \mid B_{FD} > 0\right] \cdot pr\left[B_{FD} > 0\right] \cdot S \tag{6}$$

where B_{FD} is the benefit (value) of a recreational fishing day; and S is the number of days in the season. Since the probability that the benefits are positive will be equivalent to the expected number of days of recreational fishing experienced, $E[d]$, divided by the length in days of the season, S, we have the following:

$$E\left[B_{FD} \mid B_{FD} > 0\right] \quad = \quad \left[\frac{B_L}{E[d]}\right] \tag{7}$$

Thus, we can approximate the conditional value of an expected recreational fishing day by dividing the revealed valuation of an annual license by the expected number of fishing days. This is subject to four caveats. First, these relationships assume risk neutrality. If license purchasers are risk averse, then we will over-estimate the daily valuation. This seems, however, to be a second-order problem. Second, we have ignored discounting, but it is unlikely to amount to a significant error (relative to econometric and other sources of error), considering the length of the fishing season. Third, this assumes independence of the valuation of each day, but if there is declining

[26]It is also true that an annual fishing license could offer option value for someone who decided he wanted to go fishing (for the first time in a season) in the middle of the night, when local outlets for purchasing a short-term license are closed. But this does not represent an important class of exceptions.

marginal valuation of fishing days by license holders and serial correlation among days of participation, then such independence does not hold. Fourth, depending upon the nature of unobserved heterogeneity among licensees within states, the aggregation may not produce the correct weighted average.

This takes us to the point of estimating the average expected number of fishing days per licensee per state per year, $E\,[d_{it}]$. Various approaches to this problem exist, but a reasonable approximation is simply to use the actual numbers, which are periodically compiled at the state level by the U.S. Fish and Wildlife Service (FWS).[27] This may impart a slight upward bias to the final results, since the FWS numbers refer to all (resident) participants, not only annual license holders. Some of these are short-term license holders, who likely go fishing less frequently. But the overwhelming majority of licenses sold are annual licenses, implying that this problem will not be important. A final point concerns uncertainty estimates. We use Monte Carlo procedures to simulate confidence intervals for the benefits of a recreational fishing day.

The results from this analysis are summarized in Table 5, which provides an overview of benefit estimates derived with the various specifications (all estimated as IV regressions), averaged over the sample period. Our results provide evidence of substantial heterogeneity among states in the expected value of a recreational fishing day. The mountain states plus Alaska, Arkansas, and Minnesota exhibit valuations that are ten or even twenty times the magnitude of the estimates for the lowest value states, such as Delaware, Massachusetts, and Rhode Island. This should not be surprising, and any absence of such dramatic contrasts in previous studies might even be a source of concern. Our results reflect considerable uncertainty. The 90% confidence intervals are typically large, both in absolute terms and relative to the estimated benefits of recreational fishing.[28]

In Table 6, we contrast our estimates with previous estimates of the value of a recreational fishing day, drawing upon a number of earlier studies that used either contingent valuation or travel-cost methods (but typically do not report confidence bounds).[29] Our results are generally much lower than previous estimates. Perhaps more important than the differences in absolute numbers, however, are the relative estimates of the benefits of recreational fishing. No matter which specification we use, the findings imply that fishing in states such as Colorado, Minnesota, Oregon, and Wisconsin, for example, conveys much greater recreational value than fishing in many other areas. Hence, a policymaker who were to extrapolate findings from site-specific studies of fishing in Oregon, for example, to other parts of the country could be employing a significantly biased

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[27]As reported in U.S. Fish and Wildlife Service (1975, 1980, 1985, and 1991), the surveys provide annual estimates of state-level total days of recreational fishing (separately by residents and nonresidents) and the number of participants. The estimates from those five years were interpolated to provide a set of annual estimates for the period, 1975-1989. The U.S. average over this time period is about 20 days per year for residents and 10 days per year for nonresidents.

[28]The large confidence intervals we estimate for the highest-valuation states under the log-log specifications can be misleading, however, because of the extremely asymmetric distribution of mass within them. Note the location of the expected value (mean), relative to the 90% confidence limits, for example in the case of Colorado: the confidence bounds are 13.13 and 42.03, and the mean is 13.78.

[29]We only consider states in Table 6 for which previous studies exist of the time period employed in our analysis and for which our results yield estimates that are statistically different from zero.

estimate. This would suggest particular caution when using so-called "benefit transfer" methods as frequently employed by the U.S. Environmental Protection Agency and other regulatory agencies.

Throughout the above discussion, we emphasized factors that could alter our findings. First, if there is significant option value of licenses, the actual (use value) benefits of recreational fishing would be even less than estimated here. Second, if illegal fishing increases when license prices increase, demand responsiveness will be overestimated, implying that the benefits of licenses are underestimated. Third, to the extent that other license holders may have systematically different valuations than annual license holders, our estimates will be affected. Fourth, our calculation of the expected number of fishing days relies upon USFWS statistics, which may contain errors of their own.

5. CONCLUSION

As an alternative means of estimating — on an aggregate basis — the value of a recreational fishing day, the method developed here holds promise. There is ongoing controversy surrounding the use of the contingent valuation (CV) and other hypothetical market methods for environmental-benefit estimation. Furthermore, the data requirements of the CV (and other survey methods) and travel cost models are severe, and hence the expense of carrying out such analyses is a major impediment to their use. For this reason, government agencies such as the U.S. Environmental Protection Agency rarely carry out original analyses, typically relying instead on "benefit transfers" from previous studies. Given these realities, it is of considerable value to have access to a conceptually distinct method of estimating environmental values that is based upon a revealed-preference, econometric framework.

Our numerical estimates of recreational fishing-day values suggest great variation across geographic areas, and considerably lower values than previously reported. Since previous studies have been of single sites or single states, there was inevitably some question as to whether and to what degree any observed variations were due to real differences in values, as opposed to differences among respective models. Although our approach may suffer from being a macro-oriented approach — in contrast with survey methods and travel-cost models that focus on single sites — this is also an advantage, because it facilitates the development of a set of mutually consistent estimates that can be effectively compared with one another over time and space.

TABLE 1:
MAJOR CATEGORIES OF RECREATIONAL FISHING LICENSES
CONTINENTAL UNITED STATES, 1975-1989

License Type	Total License Sales (T = 15 Years, N = 48 States)		Total Number of Observations (NT)
	Number[a]	Share	
Resident Annual	257,054,000	67.6%	720
Resident Combination[b]	52,690,000	13.9%	481
Resident Short-Term 1[c]	9,661,000	2.5%	203
Resident Short-Term 2[d]	349,000	0.1%	71
Resident Short-Term 3[e]	986,000	0.3%	29
Nonresident Annual	20,059,000	5.3%	709
Nonresident Combination	795,000	0.2%	118
Nonresident Short-Term 1	17,022,000	4.5%	378
Nonresident Short-Term 2	15,034,000	4.0%	422
Nonresident Short-Term 3	6,374,000	1.7%	203
All License Categories	380,024,000	100%	3,305

[a]Rounded to the nearest 1,000.

[b]"Combination" licenses cover both fishing and hunting on an annual basis.

[c]"Short-Term Type 1" licenses are temporary fishing licenses, ranging in length from 1 to 3 days.

[d]"Short-Term Type 2" licenses are temporary fishing licenses, ranging in length from 4 to 9 days.

[e]"Short-Term Type 3" licenses are temporary fishing licenses, ranging in length from 10 to 15 days.

TABLE 2:
RECREATIONAL FISHING LICENSES, 48 STATES, 1975-1989
DESCRIPTIVE STATISTICS

Variable	Mean[a]	Standard Deviation	Minimum	Maximum	Number of Observations
Quantity of Resident Annual Licenses per State	357,019	362,301	10,925	2,293,671	720
Quantity Per Capita of Resident Annual Licenses	0.091	0.050	0.010	0.249	720
Price of Resident Annual Licenses	$14.89[b]	3.48	$3.79	$32.80	720
Quantity of Resident Combination Licenses per State	109,543	104,386	5,059	724,990	481
Quantity Per Capita of Resident Combination Licenses	0.044	0.038	0.001	0.207	481
Price of Resident Combination Licenses	$29.02	7.84	$10.44	$72.02	481
Quantity of Resident Type 1 Short-Term Licenses per State	47,590	64,300	194	307,893	203
Quantity Per Capita of Resident Type 1 Short-Term Licenses	0.015	0.015	0.0004	0.077	203
Price of Resident Type 1 Short-Term Licenses	$9.01	3.47	$1.74	$31.77	203
Quantity of Resident Type 2 Short-Term Licenses per State	4,909	5,217	151	17,947	71
Quantity Per Capita of Resident Type 2 Short-Term Licenses	0.005	0.007	0.0001	0.026	71
Price of Resident Type 2 Short-Term Licenses	$13.43	3.36	$5.30	$22.69	71
Quantity of Resident Type 3 Short-Term Licenses per State	34,000	20,657	93	74,141	29
Quantity Per Capita of Resident Type 3 Short-Term Licenses	0.007	0.005	0.0001	0.018	29
Price of Resident Type 3 Short-Term Licenses	$6.69	1.20	$4.14	$10.44	29

[a]For states and time periods where particular types of licenses did not exist, "zero observations" have been eliminated before calculation of descriptive statistics. The unit of observation for the table is a state in a given year; there is no weighting to account for the underlying samples of unequal size.

[b]All monetary amounts throughout the paper are expressed in 2000 dollars.

TABLE 2 (CONTINUED): RECREATIONAL FISHING LICENSES, 48 STATES, 1975-1989 DESCRIPTIVE STATISTICS

Variable	Mean	Standard Deviation	Minimum	Maximu m	Number of Observations
Quantity of Nonresident Annual Licenses per State	28,292	36,570	1,232	224,850	709
Price of Nonresident Annual Licenses	$35.65	11.09	$6.40	$85.96	709
Quantity of Nonresident Combination Licenses per State	6,741	20,344	10	103,921	118
Price of Nonresident Combination Licenses	$97.52	40.10	$7.86	$250.08	118
Quantity of Nonresident Type 1 Short-Term Licenses per State	45,031	42,310	562	216,568	378
Price of Nonresident Type 1 Short-Term Licenses	$11.07	3.58	$2.64	$31.77	378
Quantity of Nonresident Type 2 Short-Term Licence's per State	35,625	38,964	661	180,712	422
Price of Nonresident Type 2 Short-Term Licenses	$16.93	5.37	$3.79	$43.33	422
Quantity of Nonresident Type 3 Short-Term Licences per State	31,398	30,568	1	185,187	203
Price of Nonresident Type 3 Short-Term Licenses	$22.19	7.26	$6.15	$47.13	203
Median Income (for four-person families)	$51,585	5,072	$36,956	$91,269	720
Share of State Population Residing in Metro. Areas	0.614	0.230	0.126	1.000	720
Mean Years of Schooling	12.67	0.40	11.54	13.54	720
State Population	4,673,888	4,911,251	377,000	29,063,00	720
State Area (in square miles)	74,377	88,748	1,212	591,004	720

Total Area of Fishable Waters (in acres)	1,845,732	4,054,958	5,983	25,416,00	720

TABLE 3:
DEMAND FOR RESIDENT ANNUAL FISHING LICENSES
FIXED EFFECTS

	Linear		**Semi-Log**		**Multiplicative (Log-Log)**	
	1	**2**	**3**	**4**	**5**	**6**
Price of Residential Annual License	-.0015171*** (.00021)	-.0015584*** (.00022)	-.0156747*** (.00195)	-.0156678*** (.00203)	-.1978754*** (.02648)	-.198246*** (.02584)
Price of Short-Term Type 1 License	.001335*** (.00035)	.0011662*** (.00035)	.006559*** (.00180)	.00678*** (.00194)	.0738751*** (.02443)	.0721838** (.02777)
Dummy/No Short-Term Type 1 License	.0139258*** (.00286)	.012949*** (.00279)	.017713*** (.00365)	.0176151*** (.00389)	.1738659*** (.04751)	.16266*** (.05245)
Price of Short-Term Type 2 License	-.0013835* (.00078)		-.009805 (.00654)		-.0187599 (.07160)	
Dummy/No Short-Term Type 2 License	-.0087098 (.00667)		-.0173988 (.01347)		.0295111 (.14443)	
Price of Short-Term Type 3 License	.0002694 (.00109)		.0037573 (.00473)		.0328461 (.07064)	
Dummy/No Short-Term Type 3 License	.0142384** (.00638)		.0173169** (.00779)		.1930563 (.12295)	
Price of Adjacent N-R Annual Licenses	-.0005014*** (.00010)		-.0164198*** (.00246)		-.1793059 (.02590)	
Quantity (Acres) of Fishable Waters	1.83×10^{-9}** (7.57×10^{-10})	3.16×10^{-9}*** (5.42×10^{-10})	.0040064 (.00268)	.0079781*** (.00240)	.042351 (.02680)	.079086*** (.02372)
Share of Population Living in Urban Areas	-.0317561* (.01614)	-.0174476 (.01761)	-.0110163* (.00660)	-.0070867 (.00695)	-.0636066 (.05285)	-.0051365 (.05332)
Median Family Income	3.69×10^{-7} (2.26×10^{-7})	5.06×10^{-7} (3.44×10^{-7})	.0165951** (.00703)	.0226339* (.01173)	.2972307*** (.07322)	.3616696*** (.10835)
Mean Years of Education	-.0075608*** (.00215)	-.0000607 (.00328)	-.1013291*** (.02270)	-.0109566 (.04199)	-.9912109*** (.24429)	-.8482652 (.60772)
Annual Fixed Effects	No	Yes	No	Yes	No	Yes
Number of Observations	720	720	720	720	720	720
R^2 (Within)	.251	.232	.274	.233	.243	.200

Note: Robust Standard Errors are in parentheses below respective parameter estimates
*** Significant at the 1% level, ** Significant at the 5% level, * Significant at the 10% level

TABLE 4:
DEMAND FOR RESIDENT ANNUAL FISHING LICENSES
INSTRUMENTAL VARIABLES

	Linear		Semi-Log		Multiplicative (Log-Log)	
Price of Resident Annual License	-.0020268 (.00061)	***	-.0394666 (.01913)	**	-.3523899 (.17869)	**
Price of Short-term Type 1 License	.001233 (.00036)	***	.0079709 (.00255)	***	.0798971 (.03264)	**
Dummy/ No Short-term Type 1 License	.0130771 (.00277)	***	.0182087 (.00468)	***	.1665044 (.05844)	***
Acres of Fishable Waters	3.45×10^{-9} (6.73×10^{-10})	***	.0100813 (.00313)	***	.0927085 (.02912)	***
Share of Population Living in Urban Areas	-.0186832 (.01800)		-.0078297 (.00700)		-.0099494 (.05406)	
Median Family Income	5.17×10^{-7} ($3.44 \times ^{10-7}$)		.0218192 (.01305)		.3563928 (.11421)	***
Mean Years of Education	-.0010463 (.00359)		-.0486377 (.06297)		-1.092323 (.72027)	
Annual Fixed Effects	Yes		Yes		Yes	
Number of Observations	720		720		720	
R-squared (within)	0.225		0.0682		0.1464	

Note: Robust Standard Errors are in parentheses below respective parameter estimates
*** Significant at the 1% level, ** Significant at the 5% level, * Significant at the 10% level

TABLE 5:
ESTIMATES OF THE VALUE OF A FRESHWATER RECREATIONAL FISHING DAY
(1975-1989 averages, in 2000 $, Based on Instrumental Variables Regressions)

	Linear IV	Semilog IV	Log-log IV cutoff $32.8	Log-log IV cutoff $100	Log-log IV cutoff $200	Log-log IV cutoff $500	90% confidence interval based on semi-log		90% confidence interval based on log-log ($200 cutoff)	
Alabama	2.63	3.86	2.61	4.23	5.72	8.52	1.14	13.12	5.08	30.84
Alaska	5.58	37.23	14.13	22.94	31.01	46.17	8.95	155.96	16.34	1901.03
Arizona	2.96	2.81	1.82	2.95	3.99	5.94	0.73	10.82	1.51	7.14
Arkansas	3.93	20.03	8.11	13.17	17.80	26.50	4.53	89.68	10.72	681.15
California	3.04	3.66	2.51	4.08	5.51	8.21	0.80	16.83	3.00	8.91
Colorado	4.60	8.79	6.28	10.20	13.78	20.52	1.81	43.20	13.13	42.03
Connecticut	1.33	1.13	0.44	0.72	0.97	1.45	0.28	4.59	0.06	3.21
Delaware	1.08	0.79	0.18	0.30	0.40	0.60	0.23	2.74	0.00	2.13
Florida	1.06	0.87	0.39	0.63	0.85	1.26	0.23	3.30	0.08	2.50
Georgia	2.20	2.28	1.66	2.69	3.63	5.41	0.70	7.51	3.44	6.09
Idaho	3.92	6.93	5.00	8.12	10.98	16.34	1.49	32.50	10.50	30.14
Illinois	1.86	1.36	0.75	1.23	1.66	2.47	0.42	4.44	0.32	3.68
Indiana	1.95	1.71	1.13	1.83	2.47	3.68	0.49	6.01	1.02	4.35
Iowa	2.84	3.42	2.49	4.04	5.46	8.13	0.90	13.11	4.96	8.57
Kansas	2.56	3.56	2.56	4.15	5.61	8.36	0.79	16.23	4.81	8.46
Kentucky	2.56	3.45	2.51	4.08	5.51	8.20	0.85	14.07	5.21	9.79
Louisiana	1.68	0.96	0.68	1.10	1.49	2.22	0.42	2.20	1.15	2.22
Maine	3.02	4.50	3.17	5.14	6.95	10.34	0.84	24.17	4.86	10.54
Maryland	1.66	1.38	0.34	0.55	0.74	1.11	0.35	5.44	0.01	3.63
Massachusetts	1.25	1.24	0.20	0.33	0.45	0.66	0.23	6.72	0.00	3.05
Michigan	2.61	3.17	2.25	3.66	4.94	7.36	0.78	12.92	4.01	7.49
Minnesota	5.02	17.98	9.81	15.93	21.53	32.05	3.73	87.01	15.28	362.81
Mississippi	1.36	0.92	0.46	0.75	1.01	1.51	0.30	2.87	0.12	2.60
Missouri	2.80	3.42	2.48	4.03	5.45	8.11	0.91	12.95	5.12	9.18
Montana	6.45	46.86	16.56	26.88	36.32	54.08	9.51	233.36	20.20	1843.60
Nebraska	3.06	4.71	3.42	5.56p	7.51	11.19	1.02	22.07	7.05	11.93
Nevada	4.03	6.10	3.94	6.40	8.65	12.88	1.06	35.45	3.66	15.47
New Hampshire	2.24	2.69	1.47	2.39	3.23	4.80	0.49	14.96	0.56	7.47
New Jersey	1.33	1.29	0.23	0.37	0.50	0.74	0.25	6.65	0.00	3.21
New Mexico	1.58	1.47	0.08	0.13	0.18	0.27	0.33	6.61	0.00	2.73
New York	1.45	1.29	0.32	0.51	0.69	1.03	0.30	5.59	0.01	3.49
North Carolina	1.27	1.18	0.27	0.43	0.59	0.88	0.25	5.54	0.01	3.25
North Dakota	4.47	8.49	5.66	9.19	12.42	18.49	2.08	34.90	10.81	69.57
Ohio	2.29	2.74	1.96	3.18	4.30	6.41	0.67	11.23	3.51	6.26
Oklahoma	2.50	3.68	2.64	4.29	5.79	8.63	0.92	14.82	5.51	16.02
Oregon	4.87	10.76	7.69	12.49	16.88	25.13	1.95	59.85	15.71	52.37
Pennsylvania	2.21	3.00	2.01	3.26	4.40	6.56	0.57	15.87	2.06	7.43
Rhode Island	0.84	0.64	0.17	0.28	0.37	0.56	0.18	2.29	0.01	1.76
South Carolina	1.58	1.48	0.85	1.38	1.86	2.77	0.37	6.07	0.41	4.02
South Dakota	3.81	5.57	3.82	6.20	8.38	12.48	1.49	20.93	7.63	26.63
Texas	2.00	1.72	1.07	1.73	2.34	3.48	0.48	6.27	0.71	4.53
Utah	4.62	8.03	5.70	9.26	12.52	18.64	1.45	44.76	10.60	19.15
Vermont	2.46	2.78	2.02	3.29	4.44	6.61	0.76	10.27	4.03	6.58
Virginia	2.01	1.94	1.20	1.95	2.63	3.92	0.48	7.86	0.76	5.09
Washington	3.09	5.61	4.09	6.65	8.98	13.38	0.99	32.13	8.12	13.57
West Virginia	2.01	2.04	1.20	1.95	2.63	3.92	0.47	8.88	0.60	5.50
Wisconsin	3.72	6.96	4.76	7.72	10.43	15.53	1.67	29.21	9.27	52.61

Wyoming	8.22	77.09	20.07	32.57	44.02	65.54	17.66	339.82	21.51	3408.97

TABLE 6:
COMPARING PREVIOUS ESTIMATES OF THE VALUE OF A FRESHWATER RECREATIONAL FISHING DAY WITH THOSE DERIVED IN THIS STUDY

State	Previous Estimates				Valuation from this Study (2000 $, based on log-log specification, cutoff $200) [90% conf.]
	Estimation Method[a]	Types of Fishing	Study	Valuation (2000 $)	
Alabama	TC	Trout	King and Hof 1985	24.57	5.86 [5.22 32.14]
Arizona	TC	All	Miller and Hay 1980	73.14	2.90 [1.09 5.14]
Colorado	CV	Cold water	Walsh, et. al. 1980	22.01	10.40 [10.01 30.38]
Georgia	TC	Warm water	Ziemer, et. al. 1980	27.65	2.92 [2.80 5.59]
Idaho	TC	All Cold/warm water	Miller and Hay Loomis and Sorg 1986	56.42 45.59/47.04	11.12 [10.80 25.18]
Maine	TC	All	Miller and Hay 1980	48.06	4.43 [2.73 6.88]
Minnesota	TC	All	Miller and Hay 1980	60.60	18.80 [13.34 305.88]
Missouri	TC	Trout	Haas & Weithman 1982	27.97	4.86 [4.62 8.06]
Ohio	TC	Cold water Perch/Walleye	Dutta 1984 Hushak, et. al. 1988	8.73 4.58/5.60	3.51 [2.60 5.24] 5.28 [4.07 7.79]
Oregon	TC	Salmon Steelhead	Brown and Shalloof 1984	36.47 49.72	12.97 [12.29 26.94]
Wisconsin	TC	All	Kealy and Bishop 1986	51.60	11.55 [10.37 52.00]

[a]CV refers to contingent valuation (survey) methods; TC refers to travel-cost (Hotelling-Clawson-Knetch) methods.

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