The Value of Nonconsumptive Wildlife Recreation in the United States

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I. INTRODUCTION AND MOTIVATION FOR ESTIMATING THE DEMAND FOR NONCONSUMPTIVE WILDLIFE RECREATION

Public interest in wildlife and wildlife-related activities extends well beyond the traditional hunting and fishing activities. In 1980, approximately 29 million individuals (16 years and older) took trips primarily to observe wildlife while 83 million (49 percent of population 16 years and older) observed specifically wildlife in both residential and nonresidential settings (USFWS 1982). Moreover, there are at least 18 national conservation or wildlife-related organizations with memberships ranging from 5,000 in the American Birding Association to 4.2 million in the National Wildlife Federation.

The extent of participation in these activities reflects a substantial demand for the products of wildlife management and yet, with few notable exceptions (Cicchetti 1973; Bishop 1978; Hay and McConnell 1979, 1984), the relationship between management programs and nonconsumptive wildlife recreation has not been investigated. Just as economic studies have contributed greatly to managers' knowledge of the value of wildlife resources for fishing and hunting (Brown, Singh, and Castle 1965; Hammack and Brown 1974; Gum and Martin 1975; Bishop and Heberlein 1979; McConnell and Strand 1981; Kealy and Bishop 1986), so they can inform on the value and extent of nonconsumptive uses of wildlife, and the sensitivity of those uses to changes in management programs.

In this paper we provide information for the wildlife management community and the public trustees by developing a model to estimate first, the probability of participation and second, the number of hours people observe, photograph, and feed wildlife away from the home conditional on participation. The model is used to show how participation in these activities may be changed with changes in habitat. In addition, we use it to calculate a lower bound estimate for the total benefits from nonconsumptive wildlife recreation trips in 1980. The data used for estimation are from the 1980 National Survey of Fishing, Hunting, and Wildlife-Associated Recreation (USFWS 1982).

II. BRIEF REVIEW OF PREVIOUS EXAMINATIONS OF NONCONSUMPTIVE USES OF WILDLIFE

Although many economists and other social scientists have studied the demand for consumptive wildlife recreation, relatively few have investigated nonconsumptive recreational demand, and these investigations have been specific rather than comprehensive. Bishop (1978) in his study of the California condor, considered the importance of nonconsumptive use, but his study was narrowly focused on that one species. Using the 1965 national survey of wildlife-associated recreation, Cicchetti (1973) estimated reduced form equations for bird watching and bird and wildlife photography. Although much broader in scope, this study is outdated. Hay and McConnell (1979) used the 1975 national survey to estimate participation equations for wildlife watching and wildlife photography, but they could not estimate the intensity of nonconsumptive wildlife activity. Hay and McConnell (1984) then estimated a joint decision model estimating participation between two recreational activities: hunting and nonconsumptive wildlife use.

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Land Economics • November 1991 • 67(4): 422–34

III. A MODEL OF NONCONSUMPTIVE RECREATIONAL USER HOURS

Suppose that the individual's total number of hours spent fishing, hunting, and observing, photographing, or feeding wildlife away from the home is separable from all other goods. Then the demand for each of these recreation activities can be estimated independently of the demand for all other goods. As in the typical travel cost demand model, we assume that the length of time the individual engages in the activity is predetermined. However, following Kealy and Bishop (1986) we allow the exogenous time at the site to vary across individuals. Then, from the individual's perspective, the choice of how many trips to take to a particular location is equivalent to the total number of hours decision.[1,2]

If we treat a nonconsumptive user hour as a generic commodity, then one equation is sufficient for estimating the demand for visitor hours at all locations assuming that all site-specific characteristics and the fishing and hunting substitutes are controlled for. In this pooled framework, the equations corresponding to the demand for hours at various locations are constrained to have identical parameters. An observation in this model is the individual's number of hours at a particular location. Therefore, individuals who visit multiple locations contribute multiple observations. For other applications of pooled travel cost demand models (i.e., varying parameter models) see Vaughan and Russell (1982) and Smith and Desvousges (1985). Also, for a favorable assessment of the performance of the pooled travel cost demand model relative to other demand models see Kling (1988). The single equation demand system

$$Z_{ij} = G(p_{zij}, p_{fij}, p_{hij}, t_{zij}, I_i, T_i, q_j, s_i) + e_{ij} \qquad [1]$$

is consistent with utility maximization if the integrability conditions are met.[3]

$Z_{ij} = D_{ij}N_{ij}$ is the ith individual's total number of nonconsumptive recreation hours spent at site j observing, photographing, and/or feeding wildlife at least one mile away from the home;

D_{ij} is the ith individual's predetermined length of stay at site j;[4]

N_{ij} is the ith individual's number of trips to site j;

$p_{zij} = p_{nij}/D_{ij}$ is the ith individual's hourly monetary cost of nonconsumptive recreation at site j including prorated expenses associated with traveling to and from the recreation site, hourly on-site costs, and prorated hourly overnight expenditures and p_{nij} is the ith individual's total monetary cost of a trip to site j for nonconsumptive recreation;

p_{fij} is the ith individual's hourly monetary cost of fishing at site j including prorated expenses associated with traveling to and from the recreation site, hourly on-site costs, and prorated hourly overnight expenditures;

p_{hij} is the ith individual's hourly monetary cost of hunting at site j constructed in the same way as for fishing and nonconsumptive recreation;

t_{zij} is the ith individual's time spent traveling to and from site j per hour of nonconsumptive recreation and one hour of on-site time;

I_i is the ith individual's income;

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[1] If we specify trips as the dependent variable, however, we lose the ability to allow time at the site to vary across individuals. Therefore, we model the number of hours decision.

[2] This formulation remains under dispute by one referee. We are still persuaded that the formulation is correct.

[3] For economic integrability to hold, the implied utility function must be quasi-concave. For the linear demand model this requires that $Z \leq -\varphi/\gamma$. Similarly, for the semi-log specification the integrability condition is $\varphi + \gamma z \leq 0$.

[4] The correlation between distance traveled and time at the site is less than 0.1. Therefore, the exogeneity assumption appears plausible for these data.

T_i is the ith individual's total discretionary time less time spent recreating at site j;

q_j is a vector of state and site quality characteristics and resource supply variables pertaining to site j;

s_i is a vector of individual i's characteristics; and

ε_{ij} is the error term due to the random component of the individual's preferences for site j.

Suppose that the consumer of nonconsumptive uses of wildlife is at an interior solution in the labor market. Then because the individual can trade time for money at the margin, she faces only one constraint, and we can rewrite [1] combining time and income terms.[5] Then,

$$Z_{ij} = G(p_{ij} + kw_i t_{ij}, p_{fij}, p_{hij}, I_i + kw_i T_i, q_j, s_i) + \varepsilon_{ij} \qquad [2]$$

or simply,

$$Z_{ij} = G(p_{zij} + kw_i t_{zij}, p_{fij}, p_{hij}, I'_i, q_j, s_i) + \varepsilon_{ij} \qquad [3]$$

where, $I'_i = I_i + kw_i T_i$ and k is a constant proportion of the individual's wage rate, w_i. We estimate two distinct specifications for [3], a linear and a semi-log to determine the sensitivity of welfare estimates to functional form. Suppressing the subscript for the individual, the linear demand equation we estimate has the form

$$Z_j = \alpha' s + \beta' q_j + \tau p_{fj} + \xi p_{hj} + \gamma I' + \varphi(p_{zj} + kwt_{zj}) + \varepsilon_j. \qquad [4]$$

We estimate equation [4] using two separate values for k, 30 and 60 percent, respectively. Similarly, the semi-log form we estimate is

$$Z_j = \exp(\alpha' s + \beta' q_j + \tau p_{fj} + \xi p_{hj} + \gamma I' + \varphi(p_{zj} + kwt_{zj}) + \varepsilon_j). \qquad [5]$$

IV. A DESCRIPTION OF THE ESTIMATION PROCEDURE AND HOW IT ACCOUNTS FOR SAMPLE SELECTION PROBLEMS

Of the previous works cited above, only Cicchetti (1973) reported estimates of frequency of nonconsumptive use, but he used ordinarly least squares (OLS) methods. If everyone in the relevant general population participated in nonconsumptive wildlife recreation, then using OLS to estimate the user hours equation would not lead to any selection bias problem. However, as has been documented in the recent recreational demand literature (Smith and Desvousges 1985; Bockstael, Hanemann, and Strand 1986; Kealy and Bishop 1986; Smith 1988) the failure to account for sample selection in the estimation of the frequency equation will bias the welfare estimates.

As with most forms of outdoor recreation, participation in nonconsumptive wildlife-related activities appeals to only a subset of the general population. In addition, it is likely that the parameters that determine participation are different from those that determine the number of nonconsumptive user hours.

There are two alternative models that allow the participation and intensity of use decisions to depend upon different factors. The first, Heckman's selection bias-corrected regression model, allows for the nonconsumptive recreation decision to be broken down into two potentially different but connected relationships: (1) the decision to participate or not to participate in nonconsumptive recreation, and (2) the choice of frequency of nonconsumptive user hours (Heckman 1979).

To obtain consistent coefficient estimates for the nonconsumptive user hours equation corresponding to the individuals who had positive consumption, first the participation equation must be estimated using a random sample from the general population. The probability of engaging in nonconsumptive recreation may be written

$$\pi = 1 - F(-X_2\beta_2/\sigma_F) \qquad [6]$$

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[5] As shown in Kealy and Bishop (1986), eq. [2] also holds for individuals at a corner of their labor/leisure trade-off if one assumes that utility from time can be expressed in monetary units as a proportion of the wage rate. In this case w_d is interpreted as the opportunity cost of leisure. See also McConnell (1975) and Bockstael et al. (1986, 4–56).

where F is the standard cumulative normal density function and the vector X includes individual and state characteristics and policy variables. The vector of parameters to be estimated is β_2, and σ_F is the standard deviation of the population's error term. Equation [6] is interesting in its own right as it is used for predicting population participation rates in nonconsumptive wildlife recreation. In addition, given the estimates of the parameters in [6], the nonconsumptive recreation user hours equations [4] and [5] can be adjusted for the bias from including only individuals with positive user hours. The selection-bias corrected regression demand equation for nonconsumptive user hours is then

$$Z_j^* = G(p_{zj} + kwt_{zj}, p_{fj}, p_{hj}, I', q_j, s) + \sigma_F\lambda_j + \xi_j. \qquad [7]$$

The estimate of the selection bias corrected variable, $\hat{\lambda}$ is obtained using the parameter estimates from [6]

$$\hat{\lambda} = F(-X_2\beta_2/\sigma_F)/[1 - F(-X_2\beta_2/\sigma_F)] \qquad [8]$$

where $F(\cdot)$ is the standard normal density function. The value for $\hat{\lambda}$ will vary across individuals depending upon the values of their respective X_2 vectors.

The second approach to modeling the two decisions, the Cragg model, is similar to the Heckman model in that it, too, uses the probit (i.e., equation [6] in the first stage) to estimate the probability of participation. Then, conditional on being a participant, in the second stage of the Cragg model the intensity of participation model is estimated independently of the first-stage probit model. This stage uses only non-zero participants (like the Heckman procedure), but the sample selection rule of the second stage in the Cragg model depends only on second-stage parameters. In this way, the Cragg model permits zero values for the level of intensity. That is,

$$Z_j^* = G(p_{zj} + kw_j t_{zj}, p_{fj}, p_{hj}, I', q_j, s) + \sigma_F\gamma_j + \xi_j \qquad [9]$$

where, $\hat{\gamma} = F(X_1\beta_1/\sigma_F)[1 - F(X_1\beta_1/\sigma_F)]$. Thus, factors that influence the participation decision can bring people in and out of participation, but also factors that influence the intensity of participation (e.g., price) can drive the level of intensity down to zero. It is this last feature that distinguishes the Cragg and Heckman models.

V. THE NATIONAL SURVEY AND HOW WE USED IT FOR VARIABLE CONSTRUCTION

A detailed description of the survey design appears in USFWS (1982). The relevant aspects of the survey include first, the telephone screening survey of U.S. households according to U.S. Bureau of the Census sampling design and second, the follow-up personal interviews of participants in nonconsumptive wildlife recreation. The screening survey interviewed about 120,000 people who were asked a variety of socioeconomic questions as well as whether anyone in the household had engaged in nonconsumptive wildlife recreation. The survey then gathered detailed trip information on about 6,000 participants in nonconsumptive activities.

From the initial phone survey, we selected randomly 2 percent (3,136 observations) of those individuals who had not engaged in taking trips to observe, photograph, or feed wildlife. We then extracted 1,155 observations from those who took at least one trip away from home to observe, feed, or photograph wildlife in the contiguous United States and who also had positive values for variable costs and time costs per hour.[6] This led to a combined working sample of 4,291 observations for estimating the participation equation [6]. For the second stage we estimated a pooled travel cost demand equation on a total of 1,769 observations. This total is greater than the number of individuals in our sample because individuals who visited multiple sites count as multiple observations. On

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[6] As we were unable to determine the mode of travel for individuals, we assumed individuals traveled to the site by automobile and calculated the time and travel costs accordingly. We excluded therefore observations of individuals who visited Alaska or Hawaii.

average, the individuals in the sample visited 1.53 sites.

The multistaged stratified sample design of the survey necessitated using a weighting procedure for estimation because the sample was, in part, choice based and nonrandom. To correct for the randomness of the sample, each observation was weighted using a sampling weight developed by the U.S. Bureau of the Census. The value of each observation's weight is the number of individuals that observation represents. A second weight was used to correct for the over-representation of nonconsumptive wildlife users in the sample. The true proportion of nonconsumptive wildlife users in the U.S. is 17 percent (USFWS 1983). For our estimation of equation [6] we employed the weighted maximum-likelihood estimator and the corrected covariance matrix as described in Manski and McFadden (1982).[7]

For the resource supply variables we have the total amount of forested acres by state and 1980 state expenditures (in thousand dollars) on nongame wildlife programs. We assigned these variables to our individual observations depending upon the individual's state of residence. For the less than 5 percent of the sample with missing data on nongame expenditures, we replaced these observations with the sample mean. A description of the variables and their construction appears in Table 1.

VI. PARTICIPATION IN NONCONSUMPTIVE WILDLIFE RECREATION

The results from estimating the PROBIT equation for participation appear in Table 2. A number of the variables are quite significant. The χ^2 statistic with 12 degrees of freedom is 2,461.2 and the pseudo R^2 is 0.42 (Judge et al. 1985, 767) indicating a reasonable goodness of fit. We compare our findings with the Hay and McConnell (1979) study who examined the probability of participation for observing wildlife and photographing wildlife using a sample from the 1975 National Wildlife Survey. They found family income divided by household size to be a statistically significant predictor of participation as did we. Neither study found the individual's sex to be a statistically significant variable. Following Hay and McConnell we included both age and age squared in our participation equation [6] and found, as they did, an inverted U-shaped participation-age relationship. They argued that including only a simple linear age variable meant that the marginal effect of a change in age on the probability of participation was constant despite the age. Constant marginal changes are improbable over the range of the variable, so we adopted their more realistic model.

One socioeconomic characteristic related negatively to participation, minority, indicates that individuals who are nonwhite are less likely to participate. Most other socioeconomic variables were statistically significant and positively related to participation. They were: the individual's years of schooling, whether the individual was married, and a retirement dummy. The employment and homemaker dummies were also positive, but they were not significant. Taken together, the results appear to indicate that income, education, and leisure time all have positive influences on participation in nonconsumptive wildlife recreation.

Of particular interest to wildlife managers are the effects of the resource supply variables. To identify the policy variables most likely to have an impact on participation, we communicated with wildlife managers at the Fish and Wildlife Service. It was they who suggested that we use total forested acres by state both because these were the best habitat data available and because woodland sites are the preferred type of site for nonconsumptive recreation.[8] Due to this preference, we hypothesized that forested acres would be related positively to participation, which is what we found.

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[7] The weights are normalized to sum to the number of observations to obtain correct asymptotic standard errors.

[8] Personal communication with Warren Fisher, Economist, USFWS.

TABLE 1

A Description of the Variables Used in the Participation and User Days Regressions

Variable	Description
INCOME	1980 gross household income divided by household size. Uses the midpoint from each of 9 income categories.
MINORITY	(0–1) variable indicating if individual was nonwhite.
EDUC	Number of an individual's years of schooling.
AGE	Age of the individual.
MARRIED	(0–1) variable indicating whether an individual was married.
EMPLOYED	(0–1) variable indicating whether an individual was employed.
HOMEMAKER	(0–1) variable indicating whether an individual was a homemaker.
National Park	(0–1) variable indicating whether site individual visited was a National park.
Woodland	(0–1) variable indicating whether site individual visited is characterized as a woodland.
Lake	(0–1) variable indicating whether site individual visited is characterized as a lake.
Hunt price	The average price per hour incurred to hunt at the site. Expenditures include the individual's cost of food, lodging, guide fees, pack trip fees, public and private land-use fees, and equipment rental fees. Mileage costs were calculated as roundtrip miles to the site times $.20/mile. These expenditures were than divided by total hours hunting at the site.
Fish price	The average price per hour incurred to fish at the site. Expenditures include the individual's cost of food, lodging, guide fees, pack trip fees, public and private land-use fees, and equipment rental fees. Mileage costs were calculated as roundtrip miles to the site times $.20/mile. These expenditures were then divided by total hours fishing at the site.
Nongame Expenditures	Amount of State Fish and Wildlife spending (in $1,000) on nongame programs in an individual's resident state.[a]
AGESQ	Age squared.
RETIRED	(0–1) variable indicating whether individual was retired.
FEMALE	(0–1) variable indicating if an individual is female.
User hours	Dependent variable in user hours equation. Total hours in 1980 to observe, photograph, or feed wildlife at least one mile away from home. Calculated as total days to each site multiplied by the average hours spent at each site.
Travel cost	The average price per day incurred to observe, photograph, or feed wildlife away from home. Expenditures include the individual's cost of food, lodging, guide fees, pack trip fees, public and private land-use fees, and equipment rental fees. Mileage costs were calculated as roundtrip miles to the site times $.20/mile. These expenditures were then divided by user hours.
Time cost	Travel time divided by average days per trip plus time spent at the site all times the wage rate. The roundtrip miles to each site were first divided by 45 miles per hour and then multiplied by the number of trips to each site divided by the total number of days to all sites. Hourly wage was calculated by dividing household income by 2,080 hours.
Total forested	Number of forested acres in each state (in acres thousands).[b]
$\hat{\lambda}$	Selection bias correction. Calculated using equation [11].

Note: All variables were obtained from the 1980 Survey of Fishing, Hunting, and Wildlife-Associated Recreation unless indicated otherwise.

[a] *Source:* USFWS 1983, 71.

[b] *Source: Statistical Abstract* 1984, 1.

The second state policy variable is the amount of State Fish and Wildlife spending on nongame programs by the state in which the individual participated in nonconsumptive wildlife recreation activities. We included this variable as a proxy for the quality of nonconsumptive user services (e.g., nature trails, guided tours, nature centers) and, although it had a positive sign, it was insignificant.

Because each of the policy variables varied only by state, they are but crude approximations for the quality of the particular location where the individual partici-

TABLE 2
PROBIT ESTIMATES OF THE PROBABILITY OF TAKING A TRIP TO PARTICIPATE IN NONCONSUMPTIVE WILDLIFE RECREATION

Variable	Coefficient	Std. Error	*t*-ratio	Mean	Std. Dev.
Income	.125E-04	.223E-10	2.651	6,828.8	5,582.7
Minority	−.579	.128E-01	−5.116	.1988	.3992
Age	.540E-01	.168E-03	4.159	33.95	21.13
Female	−.885E-01	.418E-02	−1.368	.5181	.4997
Married	.168	.539E-02	2.292	.4772	.4995
Education	.516E-01	.221E-04	10.97	10.53	5.34
Age	−.838E-03	.207E-07	−5.822	1,599	1,734.4
Retired	.663	.320E-01	3.706	.082	.2755
Employed	.151	.851E-02	1.639	.4539	.4979
Homemaker	.209	.154E-01	1.681	.1405	.3475
Total forested acres	.711E-05	.143E-10	1.880	15,466	10,419
Nongame exp.	.212E-04	.660E-09	.8230	925.9	1,844
Constant	−2.497	.749E-01	−9.121	1.00	.00
Log-Likelihood			−1,677.3		
Restricted (Slopes = 0) Log-L.			−2,907.9		
Chi-Squared (12 d.f.)		2,461.2			
Significance Level			.000		
Pseudo-*R*			.423		
Number of Observations			4,291		

Notes: Each observation was weighted by two sampling weights. The first weight, developed by the U.S. Bureau of the Census, was obtained from the screening survey. The value of each observation's weight is the number of individuals that observation represents. The second weight corrected for the over-representation of nonconsumptive wildlife users in the sample. The weighted maximum-likelihood estimator and the corrected covariance matrix was based on the technique developed by Manski and McFadden (1982).

pated in nonconsumptive recreation. Nonetheless, it is interesting to assess how the forested acres policy variable, for example, affects the probability of participation. We therefore constructed an aggregate prediction interval[9] (Amemiya 1985, 286) for current levels of the independent variables and then simulated the effect of changes in forested acres. Using the coefficients from the participation equation and the actual values of the independent variables for all observations in our sample we estimated a mean proportion of participation of 0.169599 with a standard deviation of 0.0001651. A 10 percent increase in the total forested acres by state would change the probability of participation for the sample to 0.1712276 with a standard deviation of 0.0001657. Given the U.S. population of individuals 16 years old and older of 169.924 million (USFWS 1982, 106), this change would lead to an increase of 277,000 participants with a 95 percent prediction interval of ±56,200.[10]

VII. ESTIMATION AND RESULTS OF THE USER HOURS EQUATION

Correcting for selection bias, we estimated the number of nonconsumptive wildlife recreation user hours away from the home (eq. [7]). We also corrected for heteroscedasticity using White's (1978) correction. The results for the linear and semi-log specifications appear in the first seven columns of Table 3.[11] Although we estimated two versions of each of the models, one imposing a value of time of 30 percent of the wage rate, and the other a value of 60 percent, only the former are reported in the interest of space. The sensitivity of welfare to the value of time is addressed below. We discuss the results for the semi-log specification because it is the better-fitting model. First note that most of the site characteristic variables are significant while the individual's characteristics, in general, are not significant. Travel and time cost is negative and significant.[12]

Total forested acres is positive and significant as are the lake and woodland dummies. The national park dummy also has a positive effect on visitation as does nongame expenditures, but neither variable is significant. We included price proxies for two other recreation activities, hunting and fishing, to control for the influence of substitutes and/or complements on nonconsumptive wildlife recreation visitation decisions. The price of hunting has a positive sign and the price of fishing has a negative sign, but neither are significant. Using a joint participation model, Hay and McConnell (1984) report that hunting and nonconsumptive wildlife activities are complements.

Of the individual's characteristics, only the retirement and gender dummies have a positive and significant influence on visitation. Income, age, and education are all insignificant. Finally, we note that the sample-selection coefficient, λ, is positive and significant. Both estimated equations satisfy economic integrability conditions.[13]

The results for the Cragg model, also reported in Table 3, are quite similar, but there are a few exceptions. First, age appears to have an inverted U-shape with usage first decreasing and then increasing with age. However, the retirement dummy is not significant in the Cragg model, perhaps because the influence of retirement is captured by the age variables. Also, the level of education is negative and significant in the Cragg model. Income has a negative sign, but it is not statistically different than zero. Finally, the price of hunting is positive and significant suggesting that hunting and nonconsumptive uses of wildlife may be substitutes.

To compute welfare estimates we use the closed form expressions for the appropriate Hicksian welfare measure and Marshallian consumer surplus corresponding to both demand specifications from Bockstael et al. (1986) and Arsanjani, Bockstael, Strand, and McConnell (1989). Welfare estimates for the two estimation methods (i.e., Heckman two-stage and Cragg models), the two functional forms (*i.e.*, linear and semi-log) and two values of time (i.e., 30 and 60 percent of the wage) are all reported in Table 4. Given our assumption involving the source of error in [4] and [5] (i.e., random preferences) we use the fitted values for hours of nonconsumptive wildlife recreation in our computations (Bockstael et al. 1986), but for comparison we include welfare calculations using actual hours as well.

For the linear model using time valued at 60 percent of wage, we find for our sample, the average annual willingness to pay for access to nonconsumptive wildlife recreation is 3,731 dollars per observation. From our participation model we estimate the total number of participants in the U.S. in 1980 to be 28.2 million. Multiplying the average willingness to pay per observation by our estimated number of participants and the average number of sites per participant (*1.53*) *yields a total value of the re*source of 164.5 billion 1980 dollars.

See p. 423 for "observation"

The results for the linear specification are insensitive to the source of error as-

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[9] From Amemiya (1985), the asymptotic distribution for the aggregate proportion *r*, is

$$r^A \sim N\left[n^{-1}\sum_{i=1}^{n} F_i, n^{-2}\sum_{i=1}^{n} F_i(1 - F_i)\right]$$

where *F* is the standard normal distribution.

[10] As a referee observed, we may have overestimated the responsiveness of individuals to changes in forested acres due to self-selection. People may have chosen where to live on the basis of the forest amenities offered by the state. States offering relatively few forested acres may be populated by people who are relatively less interested in forests.

[11] To avoid the problem of truncated trips using the semi-log specification, we arbitrarily selected a limit point of −5.0 rather than zero.

[12] We tried several other socioeconomic variables and resource supply variables and found that none were robust to model specification. Because the signs of these variables were sensitive to model specification we omitted them. This did not cause a problem as the *price variable was robust and significant for all* specifications.

[13] See note 3.

TABLE 3

ESTIMATION OF MODEL RESULTS

	Heckman Linear Model			Heckman Semi-log Model			Cragg Semi-log Model[b]		
(1)	(2) Coefficient	(3) Std. Error[a]	(4) t-Ratio	(5) Coefficient	(6) Std. Error[a]	(7) t-Ratio	(8) Coefficient	(9) Std. Error	(10) t-Ratio
Travel cost plus time costs at 30%	−0.56	0.103	−5.43689	−0.0132	0.00185	−7.13514	−0.0132	0.00103	−12.8155
Total forested acres	0.000348	0.000496	0.701613	0.000016	0.000004	3.761905	0.000014	0.000004	3.285372
Site is primarily woodland (0–1)	28.124	5.253	5.353893	0.42	0.0527	7.969639	0.42	0.0651	6.451613
Gross income/household size	0.000957	0.00106	0.90283	0.000002	0.000006	0.375221	−9.1E-07	0.000004	−0.20922
Age of respondent	1.404	2.667	0.526434	0.00278	0.0161	0.172671	−0.0226	0.0109	−2.07339
Age squared	−0.0235	0.0364	−0.6456	−0.00011	0.000214	−0.50467	0.00025	0.000136	1.838235
Site is primarily a lake (0–1)	18.042	6.592	2.736954	0.246	0.0565	4.353982	0.245	0.0567	4.320988
Price of hunting	2.656	2.097	1.266571	0.0511	0.0283	1.805654	0.0528	0.0258	2.046512
Price of fishing	−1.126	0.529	−2.12854	−0.0147	0.0953	−0.15425	−0.0149	0.0129	−1.15504
Respondent is retired (0–1)	53.834	46.153	1.166425	0.46	0.215	2.139535	0.257	0.178	1.44382
State non-game expenditures	0.000822	0.00222	0.37027	0.000012	0.000022	0.571429	0.000014	0.000022	0.621622
Site is a National Park (0–1)	8.548	7.177	1.191027	0.126	0.0743	1.695828	0.127	0.0716	1.773743
$\hat{\lambda}$	44.704	45.462	0.983327	0.457	0.204	2.240196			
Respondent's education	−0.485	2.382	−0.20361	−0.0166	0.012	−1.38333	−0.0321	0.00947	−3.38965
Respondent is female (0–1)	13.455	5.905	2.278577	0.109	0.0566	1.925795	0.125	0.0557	2.244165
Constant	−27.957	97.816	−0.28581	2.404	0.455	5.283516	3.279	0.238	13.77731
σ	130.214	2.189	59.48561	1.158	0.0195	59.38462	1.155	0.0194	59.53608
Log-Likelihood	−11,116			−2,762.4			−2,765		
Number of Observations	1,769			1,769			1,769		
R-squared adjusted	0.03			0.14					

[a] Corrected for heteroscedasticity following White (1978).

[b] Linear Cragg Model failed to converge.

TABLE 4

WELFARE ESTIMATES OF NONCONSUMPTIVE USES OF WILDLIFE

	Heckman Linear		Heckman Semi-log		Cragg Semi-log	
	Weighted Sample Mean	Aggregated (Millions)	Weighted Sample Mean	Aggregated (Millions)	Weighted Sample Mean	Aggregated (Millions)
Time valued at 30% of the wage						
Marshallian fitted	\$3,122 (a)	\$137,673	\$178 (d)	\$7,849	(j)	
Marshallian actual	\$2,680 (b)	\$118,182	\$1,569 (e)	\$69,189	\$1,569 (h)	\$69,189
Hicksian fitted	\$3,083 (c)	\$135,953	\$178 (f)	\$7,849	(j)	
Time valued at 60% of the wage						
Marshallian fitted	\$3,731 (a)	\$164,528	\$198 (d)	\$8,731	\$515 (g)	\$22,710
Marshallian actual	\$3,069 (b)	\$135,336	\$1,726 (e)	\$76,112	\$1,801 (h)	\$79,420
Hicksian fitted	\$3,630 (c)	\$160,075	\$198 (f)	\$8,731	\$516 (i)	\$22,754

These welfare estimates are conditional on participation. The formulas used to calculate these welfare estimates are adapted from Bockstael et al. (1986) and Arsanjani et al. (1989).

(a) $(X_0B_0)^2/-2\beta$ where X_0 is a matrix of T observations by k independent variables and B_0 is a $k \times 1$ vector of estimated coefficients (Table 3, Columns 1 and 2). β is the estimated coefficient on own price.

(b) $Z^2/-2\beta$ where Z is the $T \times 1$ dependent variable of individuals' total hours engaged in nonconsumptive wildlife activities.

(c) $-\beta/\gamma^2 + \exp(\gamma(-(X_1B_1)/\beta - p^0))[X_0B_0/\beta + \beta/\gamma^2]$ where γ is the estimated coefficient on income, X_1 is a $T \times k - 1$ matrix of observations and independent variables (excluding the price variable), and B_1 is a $k - 1 \times 1$ vector of estimated coefficients (excluding the price coefficient). p^0 is the $T \times 1$ price variable.

(d) $X_0B_0/-\beta$ (e) $Z/-\beta$ (f) $-1/\gamma\ \ln(1 + \gamma/\beta\ (X_0B_0))$ where X_0 is a $T \times k$ matrix of independent variables and B_0 is a $k \times 1$ vector of estimated coefficients (Table 3, Columns 1 and 5).

(g) $X_2B_2/-\beta$ (h) Z/β (i) $-1/\gamma\ \ln(1 + \gamma/\beta\ (X_2B_2))$ where X_2 is a $T \times k + 1$ matrix of independent variables (Table 3, Column 1) plus an added selectivity term ($T \times 1$) estimated as: $\phi(NP)/\Phi(NP)$ where N is a $T \times k$ matrix of independent variables and P is the $k \times 1$ vector of estimated coefficients (Table 3, Columns 1 and 8). ϕ is the standard normal density function and Φ is the cumulative normal. B_2 is a $k \times 1 \times 1$ vector of estimated coefficients, with the coefficient on the selectivity term the estimated σ (Table 3, Column 8).

(j) The negative coefficient on income precluded computing these values.

sumption (i.e., omitted variables or measurement error in the dependent variable). Also, the Marshallian and Hicksian welfare estimates are similar, but the different assumptions about the value of time produce differences of over 10 percent in the estimates. The largest difference in welfare estimates, however, is due to functional form. The semi-log specification estimate of $198 per observation and $8.7 billion in aggregate is substantially less than the $164.5 billion estimate for the linear model. The Cragg semi-log estimates are $515 per observation and $22.7 billion in aggregate suggesting that the estimation method (i.e., the sample selection rule) can have important effects on welfare estimates. Finally, contrary to the linear specification, for the semi-log specification the welfare estimates are sensitive to the source of error assumption. This result is most pronounced in the Heckman model where the estimates differ by an order of magnitude. Our estimates of the welfare loss due to loss of access to nonconsumptive wildlife recreation in the contiguous United States is an underestimate. The pooled travel cost demand model values loss of access to each site conditional on the unchanged characteristics of all other sites. In theory, the demand for the remaining sites should increase with each additional loss of a substitute site.[14]

VIII. SUMMARY AND CONCLUSIONS

Nonconsumptive wildlife recreation enjoys wider participation than all hunting and fishing activities combined and yet it has received comparatively little attention in the literature (USFWS 1982, 5). This study is the first to estimate the national demand for and benefits from nonconsumptive wildlife recreation using data from the 1980 Survey of Hunting, Fishing, and Wildlife-Associated Recreation. Given the limitations of the travel cost recreation demand model, we focus only on that segment of the population age 16 or older (28.8 million) who took trips at least one mile from home for the primary purpose of observing, photographing, or feeding wildlife. This compares with the total population of participants who engaged in a nonconsumptive activity for which wildlife was the primary purpose (83.2 million) (USFWS 1982, 4). Therefore, our resultant welfare measures will underestimate the total benefits derived from all forms of nonconsumptive wildlife-related activities.

We treat a nonconsumptive user hour as a generic commodity which allows us to formulate a one-equation pooled travel cost model for estimating the demand for visitor hours at each of the locations. To control for possible substitute activities, we include the prices for fishing and hunting in the demand equation for nonconsumptive wildlife recreation. A model for estimating the probability of an individual participating in a nonconsumptive wildlife activity away from home was used to account for sample selection bias using a method suggested by Heckman. An alternative method, by Cragg, was also estimated for comparison with the Heckman model. Both models were also used to assess the impact of total forested acres by state on nonconsumptive recreation decisions and to measure the effect of a potential change in this policy variable on those decisions.

Using parameter estimates from the linear user hours equation, we find an average annual willingness to pay for access to nonconsumptive wildlife recreation is 3,731 dollars per observation, but for the semi-log specification the estimate is $198. This difference in estimates due to functional form outweighed any differences due to estimation method, assumptions about the source of error or the value of time, or using Marshallian versus Hicksian welfare measures. Aggregating across all users, the total for the linear specification is $164.5 billion but for the semi-log specification the figure is $8.7 billion. Given the magnitude of participation in nonconsumptive wildlife recreation, it is important to know that the welfare estimate is sensitive to functional form. The estimates for the total net value of nonconsumptive recreation is an underestimate because the loss of access to each site was calculated assuming the existence of all other sites. For comparison, we include a table of welfare estimates for other natural resources (Table 5). Our estimates appear to lie within the range of others' welfare estimates for natural resources.

→ see p. 423 for "observation"

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[14] We owe this important observation to an anonymous reviewer.

TABLE 5

A Comparison of Natural Resource Value Estimates

Study	Region and Species Studied; Type of Model Used	Welfare Estimates
Bockstael, Hanemann, and Strand (1986)	Chesapeake Bay and Ocean City (Maryland) fishery:	
	Linear model	$2,450 to $2,564 yearly
	Box-Cox model	$2,054 to $2,135 yearly
Balkan and Kahn (1988)	U.S. deer hunting: used a linear model[a]	$1.043 per year (in 1980 dollars)
Kealy and Bishop (1986)	Lake Michigan fishery (Wisconsin side): used a linear model	$20 per day, $625 per year (in 1978 dollars)
Beurger and Kahn (1988)	New York Marine flatfish fishery: used a linear model	$968 per year (in 1986 dollars)
Bockstael, McConnell, and Strand (1988)	Nine Chesapeake Bay beaches:	
	Porter's New beach	$17 to $36 yearly
	Sandy Point	$1,080 to $2,757 yearly (in 1984 dollars)
Bockstael, Strand, and Hanemann (1987)	Southern California private boat fishing: used Tobit model	$2,703 to $4,148 yearly (in 1983 dollars)

[a] The study used the 1980 Survey of Fishing, Hunting, and Wildlife Associated Recreation.

References

Amemiya, Takeshi. 1985. *Advanced Econometrics*. Cambridge: Harvard University Press.

Arsanjani, Firuzeh, Nancy E. Bockstael, Ivar E. Strand, Jr., and Kenneth E. McConnell. 1989. "Sample Selection Bias in the Estimation of Recreation Demand Functions: An Application to Sportsfishing." Department of Agricultural Economics, University of Maryland, College Park.

Balkan, Errol, and James R. Kahn. 1988. "The Value of Changes in Deer Hunting Quality: A Travel Cost Approach." *Applied Economics* 20:533–39.

Bishop, Richard C. 1978. "Endangered Species and Uncertainty: The Economics of a Safe Minimum Standard." *American Journal of Agricultural Economics* 60 (Feb.):10–18.

Bishop, Richard C., and Thomas A. Heberlein. 1979. "Measuring Values of Extra Market Goods: Are Indirect Measures Biased?" *American Journal of Agricultural Economics* 61 (Dec.):926–30.

Bockstael, Nancy E., W. Michael Hanemann, and Ivar E. Strand. 1986. "Benefit Analysis Using Indirect or Imputed Market Methods." *Measuring the Benefits of Water Quality Improvements Using Recreation Demand Models*, Vol. 2. Washington, DC: Environmental Protection Agency, Office of Policy Analysis.

Bockstael, Nancy E., Kenneth E. McConnell, and Ivar E. Strand. 1988. "Benefits from Improvements in Chesapeake Bay Water Quality." *Benefit Analysis Using Indirect or Imputed Market Methods*, Vol. 2. EPA Contract number 811043-01. April. USEPA.

Bockstael, Nancy E., Ivar E. Strand, and W. Michael Hanemann. 1987. "Time and the Recreational Demand Model." *American Journal of Agricultural Economics* 69 (2):293–302.

Brown, William G., Ajmer Singh, and Emery N. Castle. 1965. "Net Economic Value of the Oregon Salmon-Steelhead Sport Fishery." *Journal of Wildlife Management* 29:266–79.

Buerger, Robert B., and James R. Kahn. 1988. "Conflict Between Commercial and Recreational Fishing." Prepared for the AERE Workshop on Marine and Sport Fisheries Economic Valuation and Management, June 23, Seattle.

Cicchetti, Charles J. 1973. *Forecasting Recreation in the United States*. Lexington, MA: Lexington Books.

Gum, Russell L., and William E. Martin. 1975. "Problems and Solutions in Estimating the Demand for and Value of Rural Outdoor Recreation." *American Journal of Agricultural Economics* 57 (Nov.):558–66.
Hammack, Judd, and Gardner M. Brown, Jr. 1974. *Waterfowl and Wetlands: Toward Bioeconomic Analysis*. Baltimore: Resources for the Future, Inc.
Hay, Michael J., and Kenneth E. McConnell. 1979. "An Analysis of Participation in Nonconsumptive Wildlife Recreation." *Land Economics* 55 (4):460–71.
Hay, Michael J., and Kenneth E. McConnell. 1984. "Harvesting and Nonconsumptive Wildlife Recreation Decisions." *Land Economics* 60 (4):388–96.
Heckman, James J. 1979. "Sample Selection Bias as a Specification Error." *Econometrica* 47 (1):153–61.
Judge, George G., W. E. Griffiths, R. Carter Hill, Helmut Lttkepohl, and Tsoung-Chao Lee. 1985. *The Theory and Practice of Econometrics*. New York: John Wiley and Sons.
Kealy, Mary Jo, and Richard C. Bishop. 1986. "Theoretical and Empirical Specifications Issues in Travel Cost Demand Studies." *American Journal of Agricultural Economics* 68 (3):660–67.
Kling, Catherine L. 1987. "Discrete Versus Continuous Choice Approaches to Recreational Demand Modeling." Working paper. University of California, Davis.
———. 1988. "Comparing Welfare Estimates of Environmental Quality Changes from Recreation Demand Models." *Journal of Environmental Economics and Management* 15 (Sept.):331–40.
Manski, Charles, and Daniel McFadden. 1982. "Alternative Estimates and Sample Designs for Discrete Choice Analysis." In *Structural Analysis of Discrete Data: With Econometric Applications*, eds. C. Manski and D. McFadden. Cambridge: MIT Press.
McConnell, Kenneth E. 1975. "Some Problems in Estimating the Demand for Outdoor Recreation." *American Journal of Agricultural Economics* 57 (May):330–34.
McConnell, Kenneth E., and Ivar E. Strand, Jr. 1981. "Measuring the Cost of Time in Recreation Demand Analysis: An Application to Sport Fishing." *American Journal of Agricultural Economics* 63 (Feb.):153–56.
Smith, V. Kerry. 1988. "Selection and Recreation Demand." *American Journal of Agricultural Economics* 70:29–36.
Smith, V. Kerry, and William H. Desvouges. 1985. "The Generalized Travel Cost Model and Water Quality Benefits: A Reconsideration." *Southern Economics Journal* 52:371–81.
U.S. Bureau of the Census. 1984. Statistical Abstract of the United States. 104th ed. Washington, DC: GPO.
U.S. Department of the Interior, Fish and Wildlife Service (USFWS). 1983. "A Summary of Selected Fish and Wildlife Characteristics of the 50 States." Washington, DC: GPO.
U.S. Department of the Interior, Fish and Wildlife Service (USFWS), and U.S. Bureau of the Census. 1982. "1980 National Survey of Fishing, Hunting, and Wildlife-Associated Recreation." Washington, DC: GPO.
Vaughan, William J., and Clifford S. Russell. 1982. "Valuing a Fishing Day: An Application of a Systematic Varying Parameter Model." *Land Economics* 58 (4):450–63.
White, Hal. 1980. "A Heteroskedasticity Consistent Covariance Matrix and a Direct Test for Heteroskedasticity." *Econometrica* 48:817–38.

Paying for Safety: Voluntary Reduction of Residential Radon Risks

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Radon intrusion into buildings via soil gases, water supply, and building materials is now acknowledged to be a major health hazard throughout the world. In the U.S. alone, the Environmental Protection Agency (EPA) estimates radon-associated lung cancer deaths at between 10,000 and 40,000 per year—more than any other pollutant under its jurisdiction. Because awareness of this hazard is relatively recent in the United States, there is little data on homeowner response to public appeals to test and mitigate high exposures (see Sjöberg 1989 for a recent review of existing studies).

Sweden was one of the first countries to identify radon as a public health problem, to sponsor widespread testing, and to develop reliable, low-cost techniques for reducing indoor exposure levels. Data recently collected in Sweden make it possible to derive the first empirical estimates of homeowner willingness to pay to reduce radon health risks. This paper reports these estimates and discusses their implication for public policy. We conclude that the low value of estimated willingness to pay indicates that homeowners may have interpreted the technology-based recommended exposure as a safety threshold.

After reviewing the nature and regulatory history of the radon hazard, we present a theoretical framework for modeling the value of reduced radon risk. The model employs a conventional state-dependent utility function associated with a discrete decision to mitigate or not to mitigate. In subsequent sections we describe the data collected from a sample of homeowners participating in a health department testing program and report parameter estimates from a binomial logit analysis. The estimated parameters form the basis for calculations of the probability of mitigating and willingness to pay for mitigation at various exposure levels and for various household characteristics.

I. BACKGROUND

Radon is a radioactive gas that is formed when radium decays. It is an inert gas and very little is absorbed in the human body. However, radon decay products (so-called radon daughters) are dangerous, being solid particles that can be inhaled, lodge in the lung, and emit alpha particles. Epidemiological studies of miners have attributed increased incidence of lung cancer to radon exposure.[1]

The radon protection effort in Sweden initially was directed entirely toward radiation from light concrete made of alum shale, a common building material used until 1975. It soon became apparent from measurements that most of the indoor exposures could be traced to soil gas intrusion through the foundation rather than from

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[1] There is some controversy about the implications of the miner data for indoor exposures. Nevertheless, the uncertainties are much less than for most environmental risks where the risk assessment requires extrapolating from animal studies.

Land Economics • November 1991 • 67(4): 435–46