Forest Policy Center

Internal
Working Paper Series
No. 125

Willingness to Pay for Hunting Leases in Alabama

Anwar Hussain, Daowei Zhang, and James B. Armstrong

Respectively, Research Associate, Professor and Associate Professor at the School of Forestry and Wildlife Sciences, Auburn University, AL 36849. Zhang is the corresponding author and can be reached at (334) 844-1067; Fax (334) 844-1084; E-mail: zhangdw@auburn.edu.

Willingness to Pay for Hunting Leases in Alabama

ABSTRACT: This study used a censored probit approach to estimate willingness to pay (WTP) for hunting leases in Alabama. Data were generated through a dichotomous choice contingent valuation (DCCV) survey conducted in spring 2002. After correcting for sample selection bias, we found that hunting experience, hunter's household income, number of dependents, and lack of game quality had significant impact on WTP, with lack of game quality having the largest marginal effect. The estimated WTP was $1.29/ac per hunter or $23/ac per hunting club, more than double the actual average payment of $0.52/ac per hunter or $9.36/ac per hunting club. These results suggested that landowners in Alabama could increase access fees for hunting leases.

Key Words: Hunting lease, willingness to pay, Cameron-James censored probit, dichotomous choice contingent valuation.

Willingness to Pay for Hunting Leases in Alabama

While researchers have attempted to analyze hunters' participation and associated impacts on Alabama economy (Wallace et al. 1991, Rossi 1998), the underpinnings of hunters' demand or willingness to pay (WTP) for hunting land access are not well understood. Earlier works by Stribling et al. (1992) in Alabama and Pope and Stoll (1985) in Texas used an actual hunting expense approach to analyze factors influencing hunters' participation and WTP. However, neither study came up with WTP estimates for hunting access by individuals and hunting clubs. Actual hunting expense and WTP are not the same amount unless the hunting lease market is in competitive equilibrium, and to the extent WTP is greater than actual expenditures, hunters would gain consumer surplus. Furthermore, hunting expense estimates by Wallace et al. (1991) were biased upward as their list of expenditures incurred by hunters covered the total value of vehicles, instead of the annual equivalent use value.

Goodwin et al. (1993) looked at hunters' valuation for private access in Kansas, but the study was limited as it employed an open-ended contingent valuation elicitation format. While literature on valuing non-market goods cautions about the quality of WTP estimates, it is especially critical of studies based on the open-ended contingent valuation elicitation format. This format may introduce potential biases such as strategic exaggeration of WTP or reactions to perceived hints from the survey as to appropriate valuation. Adams et al. (1989), Berrens and Adams (1998), and Fried et al. (1995) provided estimates of WTP using a censored logistic regression approach. As these studies were based on highly contrived settings, the results may not be used to make inferences in Alabama.

To our knowledge, the only estimates of WTP per year in selected groupings of states that are not actual expense estimates are provided by Boyle et al. (1998). Using a dichotomous choice contingent valuation elicitation format and censored probit regression approach, they estimated annual WTP for deer hunting at $104 per hunter. What is unknown, however, is the per acre WTP for hunting leases, which is critical for landowners who lease their lands to hunters. The purpose of this research is to estimate per acre WTP per hunter for hunting leases in Alabama. Results of this study may be useful to Alabama landowners in considering how to adjust hunting lease rates. The methodology can also be useful in studying hunting leases in other regions of the United States.

Determinants of Hunting Lease Purchase and WTP

The decision to purchase a hunting lease and whether or not you would be willing to pay more as a lessee are distinct but interrelated decisions. Economic theory suggests that demand for a commodity depends on its own price, prices of substitutes, income, and socio-demographics. In this context, factors influencing demand for a hunting lease were thus hypothesized to include hunting lease rate, alternative hunting access options, household income adjusted for number of dependents, investment in hunting equipment, hunting experience, location of residence, and hunting club membership. All else equal, the lower the lease rate, the higher the probability that a hunting lease would be purchased, and in the extreme circumstance where a hunter had free access to a hunting site, we would expect him/her not to buy a lease.

Furthermore, factors influencing the WTP of a lessee could be identified as: (1) hunter's characteristics such as hunting experience, income, and residence location; (2) game characteristics such as harvest success, diversity, and quality; (3) hunting site characteristics such as relative concentration of hunters, accessibility, and overall quality of setting; and (4) pre-determined bid price. In particular, we would expect that increases in household income adjusted for number of dependents, harvest success, and site quality to increase WTP and that hunting experience, lack of game quality, and bid price to have the opposite effect. Of these factors, the status of harvest success in explaining hunters' participation and satisfaction, with associated implications for the design of wildlife management strategies, has been a subject of particular interest (Hayslette et al. 2001). The factors used in this study along with their hypothesized effects on the probability of hunting lease purchase and the WTP of a lessee are summarized in Table 1.

Methods

WTP Estimation and DCCV

Estimating willingness to pay for non-market goods requires two steps—data generation and estimation. The dichotomous choice contingent valuation (DCCV) approach pioneered by Bishop and Heberlein (1979) has often been used as the mechanism for generating data. The DCCV approach first involves establishing attributes of the non-market good in question, and then asking the respondent whether or not he/she would pay or accept a single specific offer or bid price to access the resource or good. The respondent merely decides whether to accept or refuse the offer. With DCCV, researchers do not know the exact magnitude of the respondent's valuation, but only know whether it is greater than or less than some specified amount predetermined by researchers. The arbitrarily assigned sums vary across respondents. This strategy is attractive as it generates a scenario that is similar to day-to-day market transactions for each respondent. Furthermore, it is less stressful for respondents to say "yes" or "no" rather than to require them to give a specific value (Cameron and James 1987). Thus, this approach circumvents much of the potential bias due to strategic responses. The drawback is that WTP must be inferred, and the resulting estimate may be sensitive to the assumptions about utility function, distribution of error term, and associated functional form (Loomis 1990).

Hanemann (1984) and Cameron and James (1987) provided justification for use of dichotomous contingent valuation based on utility theory and have developed their respective WTP estimation methods. While the Hanemann approach required explicit specification of utility functions and derived functional forms for WTP estimation, the Cameron-James approach specifies a logistic or probit model for estimation that can be transformed into a valuation function. Park et al. (1991) provided the first empirical test of the connections between the Hanemann and Cameron-James approaches and found that welfare estimates under both were not significantly different. McConnell (1990) showed that when the marginal utility of income was constant, the models under both approaches were linear transformations of one another. However, the Cameron-James approach has become popular because it is easy to implement and interpret. In this study, we used the Cameron-James estimation approach, with a DCCV as the data generating mechanism.

Cameron-James Censored Probit Approach

The Cameron-James approach specifies a WTP function based on a censored probit or censored logistic regression. The approach starts with a conventional binary probit regression which includes the bid price (t_i) as an explanatory variable, and the WTP function is derived from it by dividing the explanatory variable coefficients by the negative coefficient (α) of the bid price t_i. The WTP function can assume linear or log-linear functional forms, meaning that the bid price can be entered in linear or log-linear form. Because the Cameron-James approach allows WTP to be modeled as a linear function of the explanatory variables, the estimated coefficients can be interpreted as the change in WTP for a unit change in any of the explanatory variables, an interpretation that cannot be made of conventional probit coefficients. Predicted WTP of lessee i is then obtained by inserting his/her values of the explanatory variables into the fitted WTP function. Mean ($\overline{WTP}$) estimates for linear and log-linear functional forms, given a sample of *m* lessee hunters, would respectively be given by:

$$\overline{WTP} = \sum_{i=1}^{m} \hat{WTP}_i)/ m \qquad (1a)$$

$$\overline{WTP} = (e^{(\sum_{i=1}^{m} \ln \hat{WTP}_i)}) / m \qquad (1b)$$

where $\hat{WTP}_i$ is the predicted WTP associated with lessee *i*.

Note that judicious selection of the bid price t_i and bid set from which t_i is drawn can result in large gains in the efficiency of the estimates, reducing the need for large sample sizes often required with binary data. In this regard, various proposals have been advanced. Cooper and Loomis (1992) suggested covering as much of the WTP range as possible to minimize the biases of the WTP estimates. Kanninen (1995) showed that a more complete WTP range had only minimal effects on WTP estimates. Alberini (1995) suggested using four to six bids and avoiding bids in the extreme tails of WTP. Finally, Arrow et al. (1993) recommended that respondents be reminded of substitutes as well as their limited budget while eliciting WTP. In this study, we used 10 prices in the bid set with $1 increment.

Sample Selection Bias

Since some hunters did not lease, a straightforward implementation of Cameron-James approach using univariate probit on those who leased may not yield consistent estimates as it introduces a sample selection bias (by dropping those who did not lease). We used a bivariate probit model to account for sample selection bias (Greene 1993, Whitehead et al. 1993). Given a sample of *N* hunters, we have

$$y_1^* = x_1\beta + \varepsilon_1 \qquad (2a)$$

$$y_2^* = x_2\gamma + \varepsilon_2 \qquad (2b)$$

$$E[\varepsilon_1] = E[\varepsilon_2] = 0;\ var[\varepsilon_1] = var[\varepsilon_2] = 1;\ cov[\varepsilon_1, \varepsilon_2] = \rho$$

The unobservable variables y_1^* and y_2^* are respectively related to the binary (observed) outcomes by the selection and outcome equations:

$$y_1 = x_1\beta + \varepsilon_1 \qquad y_1 = 1,\ y_1^* > 0, \text{ i.e., if hunter is lessee;}$$
$$0, \text{ otherwise} \qquad (3a)$$

$$y_2 = x_2\gamma + \varepsilon_2 \qquad y_2 = 1, y_2^* > 0, \text{ i.e., if lessee is willing to pay the bid price;}$$
$$0, \text{ otherwise} \qquad (3b)$$

where β and γ are vectors of parameters associated with the set of variables in x_1 and x_2 that need to be estimated. A test of the hypothesis of no sample selection bias is provided by H_o: $\rho = 0$. Failure to reject this hypothesis means that WTP estimates based on univariate probit regression, which uses observations on lessees only and constrains the correlation between the error terms in the selection and outcome equation to zero, could be used to make inferences regarding all hunters, lessees as well as non-lessees.

Data

Survey Implementation

To implement the DCCV, a pre-tested survey questionnaire, along with cover letter and a $1 bill was mailed to 622 active hunters in the spring 2002.[1] Following a series of questions as

[1] The list of active hunters was obtained from a two phase survey conducted in spring 2001. Active hunters were defined as those who had hunted at least once in three years prior to 2001. See Zhang and Armstrong (2002) for a detailed description of the two phase survey.

to whether or not a lease was purchased, lessee hunters were asked to respond to the following dichotomous choice contingent valuation question: *"Would you have leased the same land you hunted on, had the lease rate per acre been higher by* $\$t_i$*."* The bid price ($\$t_i$) ranged from \$1 to \$10, in increments of \$1 and randomly varied across respondents. To remind respondents of their limited budgets, the dichotomous choice contingent valuation question was preceded by series of questions on access options and their hunting trip related expenses.

Nineteen unopened surveys were returned due to wrong addresses, reducing the sample size to 603. A total of 335 questionnaires were returned after two mailings, representing a response rate of 56%. Because a relatively high response rate was achieved, a non-response check was not conducted. Of the 335 responses, 20 responses were not usable due to incomplete information on various variables of interest. The remaining 315 respondents comprised of 121 lessees and 194 non-lessees. As the predetermined bid price $\$t_i$ was addressed to lessees, it was essentially this set of observations that was used to estimate WTP for hunting leases while accounting for sample selection bias by using the bivariate probit framework.

The majority (107 or 88%) of the 121 lessees were members of a hunting club. The average number of hunters per hunting club was 18. However, not all of these 107 club members purchased hunting leases as club members. Rather 100 (83%) lessees purchased hunting leases as members of a hunting club—leases that were mainly all game-annual and cost less per hunter than leases purchased independently by the remaining 21 (17%) lessees.[2] In addition to spending an average sum of \$53 per trip on gasoline and food, respondents on

[2] It was not clear why the remaining 7 responding hunters did not purchase hunting leases as club members.

average leased a total of 1,582 ac for a total cost of \$827/ hunter, or \$0.52/ac per hunter. Given the average hunting club size of 18 recorded in this study, this translated to an average lease rate of \$9.36/ac/year.

Of the 121 lessees, 39 were willing to pay the pre-set additional amount of money to retain their hunting leases, and 82 were not. Comparing the 39 hunters who were willing to pay the higher rate with the 82 hunters who indicated they were not, we noticed that a relatively higher percentage of the latter category expressed dissatisfaction regarding quality, abundance, and diversity of game. A sub-sample of 53 lessees (44% of 121) who reported having other land access options such as National Forests still purchased hunting leases—suggesting that some hunters prefer private hunting lands over public and other types of hunting lands. This could be because private hunting lands have less hunter congestion, offer better harvest success, and are in large lot sizes. The hunting sites covered in this study were 57% forests, 24% agriculture, and 19% other land use types (i.e., swamps, wetlands, creeks, and rivers). Descriptive statistics of variables used in this study are presented in Table 2.

Variables Construction

Hunters generally lease hunting land to experience a range of satisfactions rather than just harvesting particular species (Gigliotti 2000) and any measure of harvest success should have ideally considered harvest of all species and other satisfactions realized. Constructing such an index of success was, however, fraught with problems such as the appropriate weights for different aspects of the hunting experience. Thus, we constructed an index that involved only successful white-tailed deer (*Odocoileus virginianus*) and eastern turkey (*Meleagris galloparo silvestris*) harvest per trip because these are the species most often pursued and harvested in Alabama (U.S. Fish and Wildlife Service, 1998). Finding a proxy for site quality was equally challenging. Following Clark and Stankey (1979), alternative sites were characterized as primitive and semi-modern in the recreation opportunity spectrum framework. Finally, to account for hunter avidity, a dummy variable was constructed based on the criterion that those who had invested over $5,000 in hunting equipment and vehicles were placed in a separate category than those who had invested less.

Results and Discussion

Testing for Sample Selection Bias

Before estimating WTP for hunting leases, we used a bivariate probit regression to test for sample selection bias if observations on non-lessees were ignored. Results of the probit with sample selection, estimated using full information maximum likelihood (FIML), are given in Table 3. In the selection equation, which predicted the probability as to whether a hunter would purchase a hunting lease or not, the dependent variable, LSENLSE, is equal to 1 if a hunter purchased a hunting lease and 0 otherwise. Independent variables included hunting experience, household income, number of dependents, hunting club membership, investment in hunting equipment, location of residence, and alternative hunting access options. In the outcome equation, which related the probability as to whether a lessee would be willing to pay the pre-specified additional amount or not, the dependent variable, WTP, is equal to 1 if the lessee was willing to pay the bid price and 0 otherwise. The independent variables in the outcome equation included the natural logarithm of the specified amount to be paid, hunting experience, household income, number of dependents, average harvest success per trip, lack of game quality, and site quality.

The rho (ρ) statistic is the correlation between error terms in the selection and outcome equation. Selection bias is then tested for by using the likelihood ratio test. Since the value of estimated log likelihood ratio statistic with 1 degree of freedom was 0.11 and the value of χ^2 with one degree of freedom at a 5% level of significance was 3.84, the null hypothesis of no sample selection bias could not be rejected. This implied that WTP estimates based on the univariate probit regression were valid for making a reference to the whole study sample. Comparing the univariate probit results with the bivariate probit results showed that the same coefficients had the same sign. The coefficient on harvest success, however, was significant in the univariate probit model but only marginally significant in the bivariate probit model.

Estimation of WTP

Table 4 presents the results of the bivariate probit model that were used to impute the WTP for hunting leases in Alabama. All estimated coefficients had signs consistent with expectations and, with the exception of the coefficients on harvest success and site quality variable, all other coefficients were statistically significant at the 10% level or better. Thus, increases in household income were more likely to increase the probability that a lessee would be willing to pay the proposed extra amount while increases in hunting experience, perceived lack of quality game, and number of dependents reduced the probability. To get a better perspective on the response of WTP to changes in the explanatory variables, marginal effects evaluated at the means of explanatory variables are given. Thus, a unit increase in household income increased the probability of paying the proposed extra amount by 0.3%, and lack of quality game reduced the probability of paying the proposed extra amount by about 17%.

WTP estimates were derived based on the Cameron-James approach, and the estimated results are also presented in Table 4. These estimates can be interpreted like standard regression coefficients. For example, as household income increases by $1,000, WTP increases by

$0.002/ac. Likewise a unit increase in perceived lack of game quality reduces WTP by $1.24. Using equation (1b), the overall average WTP was estimated as $1.29/ac per hunter.

To provide a perspective on the relative size of our estimates, we compared our results with two types of WTP estimates in the literature—WTP for private access to an unspecified number of acres and WTP for some probability of harvesting a deer. For the first type, Goodwin et al. (1993) estimated that WTP for private hunting land access in Kansas was $81.23 per year. Stribling et al. (1988) found respondents willing to pay an additional $0.16/ac if they could harvest an additional deer and about $0.42/ac more if they could harvest three or four more deer. Although not directly comparable to our results, these numbers seems to be lower than those found in this study.

As to the second type of WTP estimates, Livengood (1983) showed that hunters would be willing to pay about $25, on average, to be assured of harvesting one deer and $13 for a second deer. Mackenzie (1990) suggested that hunters in Delaware were willing to pay a marginal value of $6.84 for a 1% increase in the probability of harvesting a deer. Similarly, Fried et al. (1995) estimated that hunters would be willing to pay $287 for a virtually certain opportunity over a 4-day period to shoot at an Elk (*Cervus Elaphus*) at the Starkey Forest, Oregon. For selected groupings of states, Boyle et al. (1996) estimated the average net WTP per year for deer hunting at $104.

Finally, this study did not find strong empirical support for the centrality of harvest success to hunter satisfaction. This finding is consistent with Hayslette et al. (2001) who found that most Alabama hunters appeared motivated by multiple satisfactions.

Conclusions

This study estimated that the demand for hunting leases or WTP per hunter was \$1.29/ac (equivalently \$23.22/ac per hunting club, given an average hunting club membership of 18) as opposed to an actual average payment of \$0.52/ac per hunter or \$9.36/ac per hunting club. Since existing hunting lease fees in Alabama ranged from \$5 to \$12 per acre, these results suggested a potential for increasing lease fees by landowners. Given that stumpage prices were depressed recently in Alabama, an increase in hunting lease fees could enhance landowner income as well as wildlife management activities on their lands. However, since hunting sites covered in this study are usually large, these results may not be applicable to landowners with small or medium acreage. Since large tracts are fewer and scarcer than small and medium tracts, owners of large tract could get a higher hunting lease fees for their tracts.

However, there is a difference between willingness-to-pay and actual hunting fees for hunting leases. Willingness-to-pay is a measure of demand expressed by hunters, and actual hunting fees are specific points on demand curve. Whether and how much landowners can "squeeze" more from hunters and maximize their hunting lease fess depend on many factors—the bargaining powers of landowners and of hunters, and the characteristics and location of the site as well as these of substituting sites. Thus, a higher (than actual hunting fee paid) willingness-to-pay estimate, as shown in this paper, does not mean landowners can easily extract a higher hunting fee. Rather this paper only shows that the potential for some large landowners to increase hunting fees exists. Furthermore, given that the demand for hunting leases is more likely to be elastic, possibly due to substitute hunting sites and the view that hunting is a luxury good, lease rates need to be increased only in small increments overtime.

More importantly, landowners and consulting foresters, who may advise their clients on hunting leases, should set up a lease fee that reflects the number of hunters or the size of hunting clubs. Large landowners could enhance their site and game quality and market their forestlands to larger hunting clubs. Similarly, these hunting clubs could use their size to reduce the per-capita hunting fees when leasing lands.

In regard to the relative importance of various factors in explaining WTP, this study found that hunting experience, household income, number of dependents, pre-determined bid price, and lack of game quality were statistically significant and that site quality and harvest success were marginally significant. Given the exogenous nature of hunter characteristics such as experience and household income, forest landowners could increase hunting fees by focusing on enhancement of game and site quality on their lands.

Table 1. List of variables and hypothesized effect on the WTP for a hunting lease.

	Definition	Expected sign
SELECTION EQUATION	Probability of being a lessee	
Dependent variable	LSENLSE - Dichotomous: 1 if lessee; 0 otherwise	
Explanatory variable		
Hunting experience	Years of hunting	-
Income	Household income in $1,000	+
Dependents	Number of dependents in hunter's family	-
Hunting club membership	Dummy variable: 1 if yes; 0 otherwise	+
Investment in hunting equipment	Dummy variable: 1 if value of hunting equipment at least $5,000; 0 otherwise	+
Residence	Dummy variable: 1 if urban; 0 otherwise	+
Substituting hunting access option	Dummy variable: 1 if exists, 0 otherwise	-
OUTCOME EQUATION	Probability of paying the bid price conditional on being a lessee	
Dependent variable	WTP-Dichotomous: 1 if willing to pay the proposed extra amount per acre; 0 otherwise	
Explanatory variable		
Hunter characteristics		
Hunting experience	Years of hunting	-
Income	Household income in $1,000	+
Dependents	Number of dependents in hunter's family	–
Game characteristics		
Hunting success	Average number of deer and turkey harvested per trip	+
Lack of quality game	Dummy variable: 1 if hunter expressed dissatisfaction with game quality; 0 otherwise	–
Site characteristics		
Site quality	Dummy variable: 1 if perceived site is primitive; 0 if semi-modern	-/+
Predetermined bid ($)	Randomly varied across respondents in the interval {1 to 10 dollars} with bid sets of 1, 2, 3, …..10.	–

Expected sign: (-) an increase in the explanatory variable is hypothesized to decrease the dependent variable; (+) an increase in the explanatory variable is hypothesized to increase the dependent variable; (-/+) hypothesized effect could take either sign depending on hunter preferences.

Table 2. Mean and standard deviation of the variables used in estimation.

VARIABLE	All hunters (n =315)		Lessee hunters (n =121)	
	Mean	SD	Mean	SD
Hunter characteristics				
Hunting experience	32.38	15.11	30.88	14.98
Income	64.06	34.03	72.19	35.71
Dependents	1.56	1.35	1.61	1.39
Hunting club membership	0.39	0.49	0.88	0.32
Investment in hunting equipment	0.26	0.44	0.42	0.50
Residence	0.38	0.49	0.41	0.49
Substitute hunting access option	0.34	0.48	0.23	0.42
Game characteristics				
Hunting success	0.21	0.34	0.19	0.27
Lack of quality game	0.41	0.49	0.48	0.50
Site characteristics				
Site quality	0.39	0.49	0.40	0.49
Predetermined bid ($)	2.11	3.20	5.50	2.85
Natural log (predetermined bid)	0.59	0.84	1.53	0.64

Table 3. Maximum likelihood estimates of univariate probit and full information maximum likelihood (FIML) estimates of a bivariate probit regression.

	Univariate probit		Bivariate probit	
	Coefficient	t-ratio	Coefficient	t-ratio
SELECTION EQUATION	Probability of being a lessee(n =315)			
Constant (β_0)	-1.803***	-4.986	-1.798***	-4.655
Hunter characteristics				
Hunting experience (β_1)	-0.011*	-1.582	-0.011*	-1.579
Income (β_2)	0.006**	1.878	0.006**	1.813
Dependents (β_3)	-0.024	-0.291	-0.023	-0.221
Hunting club membership(β_4)	2.683***	11.860	2.685***	11.361
Investment in hunting equipment (β_5)	0.531***	2.187	0.531**	1.887
Residence (β_6)	0.248	1.062	0.243	0.857
Substitute hunting access option (β_7)	0.134	0.577	0.131	0.509
Log-likelihood	-87.99			
OUTCOME EQUATION	Probability of paying the bid price conditional on being a lessee (n =121)			
Constant (γ_0)	0.295	0.560	0.303	0.530
Hunter characteristics				
Hunting experience (γ_1)	-0.019***	-2.189	-0.019***	-1.942
Income (γ_2)	0.009***	2.400	0.009***	2.314
Dependents (γ_3)	-0.212***	-2.069	-0.210*8	-1.757
Game characteristics				
Harvest success (γ_4)	0.863**	1.770	0.875*	1.432
Lack of quality game (γ_5)	-0.521***	-1.914	-0.512**	-1.784
Site characteristics				
Site quality (γ_6)	0.423*	1.509	0.444*	1.358
Log predetermined bid (α)	-0.411***	-1.924	-0.411**	- 1.677
Rho (ρ)	0.0		-0.087	-0.267
Log-likelihood	-62.66		-150.59	

*** Significant at the 5% level.
** Significant at the 10% level.
* Significant at the 20% level.
Likelihood ratio: $\chi^2(1) = 0.11$

Table 4. Full information maximum likelihood (FIML) estimates of a bivariate probit regression, marginal effects, and WTP function.

Variable	Coefficient(γ_j)	t-ratio	Marginal effect†	Implied WTP function: $-\gamma_j/\alpha$
Constant (γ_0)	0.303	0.530	0.062	0.738
Hunter characteristics				
Hunting experience (γ_1)	-0.019***	-1.942	-0.007	-0.047
Income, thousand dollars (γ_2)	0.009***	2.314	0.003	0.002
Dependents (γ_3)	-0.210**	-1.757	-0.070	-0.511
Game characteristics				
Harvest success (γ_4)	0.875*	1.432	0.288	2.129
Lack of quality game (γ_5)	-0.512**	-1.78	-0.168	-1.243
Site characteristics				
Site quality (γ_6)	0.444*	1.358	0.146	1.081
Log predetermined bid (α)	-0.411**	- 1.677	-0.135	
Log-likelihood function	-150.59			
Estimated WTP/hunter/acre ($)				1.29

*** Significant at the 5% level.
** Significant at the 10% level.
* Significant at the 20% level.
† Computed at the means of the explanatory variables. For dummy variables, the change in probability is calculated due to the change in the value of independent variable from 0 to 1.

Literature Cited

Adams, R.M., O. Bergland, W.N. Musser, S.L. Johnson, and L.M. Musser. 1989. Use fees and equity issues in public hunting expenditures: the case of ring-necked pheasant in Oregon. Land Econ. 65 (4):376-385.

Alberini, A. 1995. Testing willingness to pay models of discrete choice contingency valuation survey data. Land Econ. 7(1):83-95.

Arrow, K., R. Solow, P. Portney, E. Learner, R. Rodner, and H. Schuman. 1993. Report of the NOAA panel on contingent valuation. Federal Register 58(10):602-614.

Berrens, R.P., and R.M. Adams. 1998. Applying contingent valuation in the design of fee hunting programs: pheasant hunting in Oregon revisited. Human Dimen. Wildl. 3(3):11-25.

Bishop, R.C., and T. Heberlein. 1979. Measuring values of extra-market goods: are indirect measures biased? Amer. J. Agr. Econ. 61(5):926-930.

Boyle, K.J., B. Roach, and D.G. Waddington. 1998. 1996 Net economic values for bass, trout and walleye fishing, deer, elk and moose hunting, and wildlife watching: Addendum to the 1996 national survey of fishing, hunting and wildlife associated recreation. Report 96-2. U.S. Fish and Wildlife Service. Washington, D.C. 20240.

Cameron, T.A., and M.D. James. 1987. Efficient estimation methods for close-ended contingent valuation surveys. Rev. Econ. Stat. 69(2):269-76.

Cameron, T.A. 1988. A new paradigm for valuing non-market goods using referendum data: maximum likelihood estimation by censored logistic regression. J. Environ. Econ. Manage. 15(3):355-379.

Clark, R., and G. Stankey. 1979. The recreation opportunity spectrum: a framework for planning, management, and research. USDA For. Serv. Research Paper PNW-98. Portland, Oregon 97232.

Cooper, J., and J.B. Loomis. 1992. Sensitivity of willingness to pay estimates to bid design in dichotomous choice contingent valuation models. Land Econ. 68(2):211-24.

Fried B.M., R.M. Adams, R.P. Berrens, and O. Bergland. 1995. Willingness to pay for a change in elk hunting quality. Wildl. Soc. Bull. 23(4):680-686.

Gigliotti, L.M. 2000. A classification scheme to better understand satisfaction of Black Hills deer hunters: the role of harvest success. Human Dimen. Wildl. 5(1):32-51.

Goodwin, B.K., L.A. Offenback, T.T. Cable, and P.S. Cook. 1993. Discrete/ continuous valuation of private hunting access in Kansas. J. Environ. Manage. 39(1):1-12.

Greene, W.H. 1993. Econometric analysis. 2nd edition. Macmillan Publishing Co, New York, NY.

Hanemann, W.M. 1984. Welfare evaluations in contingent valuation experiments with discrete choice responses. Amer. J. Agr. Econ. 66(3):332-341.

Kanninen, B.J. 1995. Bias in discrete response contingent valuation. J. Environ. Econ. Manage. 28(1):114-125.

Hayslette, S.E., J.B.Armstrong, and R.E. Mirarchi. 2001. Mourning dove hunting in Alabama: motivations, satisfactions, and sociocultural influences. Human Dimen.Wildl. 6(2):81-95.

Livengood, K.R. 1983. Value of big game from markets for hunting leases: the hedonic approach. Land Econ. 59(3):287-291.

Loomis, J.B. 1990. Comparative reliability of the dichotomous choice and open-ended contingent valuation techniques. J. Environ. Econ. Manage. 18(1): 78-85.

Mackenzie, J. 1990. Conjoint analysis of deer hunting. Northeast. J. Agr. Resource Econ. 9(2): 109-117.

McConnell, K.E. 1990. Models for referendum data: the structure of discrete choice models for contingent valuation. J. Environ. Econ. Manage. 18(1):19-34.

Park, T., J.B. Loomis, and M. Creel. 1991. Confidence intervals for evaluating benefits estimates from dichotomous choice contingent valuation studies. Land Econ. 67(1):64-73.

Pope III, C.A., and J.R. Stoll. 1985. The market value of ingress rights for white-tailed deer hunting in Texas. South. J. Agr. Econ. 17(1):177-182.

Rossi, A. N. 1998. Alabamians' hunting behavior, wildlife-related activities, and wildlife management perceptions. MS Thesis, Auburn Univ., AL.

Stribling, H.L., J.P. Caulfield, B.G. Lockaby, D.P. Thompson, H.E. Quicke, and H.A. Clonts. 1992. Factors influencing willingness to pay for deer hunting in the Alabama piedmont. South. J. Appl. For. 16(3):125-129.

U.S. Fish and Wildlife Service. 1998. 1996 National survey of fishing, hunting, and wildlife-associated recreation. FHW/96-AL. USGPO, Washington, D.C.

Wallace, M. S., H. L. Stribling, and H.A. Clonts. 1991. Effect of hunter expenditure distribution on community economies. Wildl. Soc. Bull. 19(1):7-14.

Whitehead, J.C., P.A. Groothuis, and G.C. Bloomquist. 1993. Testing for non-response and sample selection bias in contingent valuation: analysis of a combination phone and mail survey. Econ. Letters. 41(2):215-220.

Zhang, D., and J.B. Armstrong. 2002. Causes of declining hunting license sales in Alabama. A Report to the Alabama Department of Conservation and Natural Resources. School of Forestry and Wildlife Sciences, Auburn Univ., AL.