# Analyzing Agricultural Landowners' Willingness to Install Streamside Buffers

*by*

Lori Lynch, Ian Hardie, and Doug Parker

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Department of Agricultural and Resource Economics

The University of Maryland, College Park

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Department of Agricultural and Resource Economics
2200 Symons Hall
University of Maryland
College Park, Maryland, 20742 USA
(301) 405-1264
(301) 314-9091 (fax)
llynch@arec.umd.edu

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**Abstract:**

Many watershed organizations have prioritized establishing streamside (riparian) buffers on agricultural land to improve water quality. Using data from a 2000 survey of 500 Maryland landowners, we examine what level of financial incentives they would require to install such buffers for 15 years on a voluntary basis. A random utility model is developed where a landowner is willing to accept the offered contract if he or she receives a higher utility from the incentive payment and buffer installation than from not planting the buffer. Given the development pressure in the Washington D.C./Baltimore corridor, we test whether farmers need more than the agricultural opportunity costs to encumber their land. Higher incentive payments, part-time farming, education, and a Lower Shore location positively influence the respondent's willingness to install a buffer. Length of the farming horizon, age, and a Southern Maryland location negatively influence the respondent's willingness.

## Introduction

Economic analysis of the use of streamside (riparian) buffers and filter strips for water quality improvement has been increasing (Kline, Alig and Johnson; Lynch and Brown; Lohr and Park; Purvis, et al.; Olmstead and McCurdy), but does not yet approach the more than 400 papers that have been published describing the buffer's physical and chemical processes since 1970 (Correll). Buffers have been found to improve water quality by filtering, transforming, and absorbing agricultural nutrient runoff, removing 50-85% of the nitrogen and phosphorus, as well as other agricultural chemicals, from the water before it enters the stream. Such findings have encouraged watershed organizations to give a high priority to establishing riparian buffers. The Chesapeake Bay Program, for example, has set a goal to install forest riparian buffers on 2,010 miles of streams by the year 2010 (1.8% of the stream miles). Vermont has established land-use regulations that seek to improve water quality through streamside buffers (Sanford and Stroud). Oregon is encouraging landowners to install streamside buffers to decrease stream temperature and improve fish habitat (Mooney and Eisgruber). Yet, agricultural landowners' willingness to install these buffers has not been studied.

Buffer programs may encourage landowners to voluntarily install or maintain buffers along their water bodies. These programs often provide financial incentives to install buffers to compensate landowners for taking land out of production. One such program enhances the Conservation Reserve Program (CRP) to target specific environmental goals such as water quality and wildlife habitat. In 1997, the U.S. Department of Agriculture (USDA) in cooperation with Maryland state agencies and nonprofit groups introduced the Conservation Reserve Enhancement Program (CREP). This voluntary incentive program provides annual rental payments, incentive bonuses and cost-share payments to

Maryland agricultural landowners who install buffers.  As of December 2001, CREP had enrolled more than 27,800 acres of grass and forest buffers, about 28 percent of the 100,000 acre buffer goal established for the program.

We employ  a stated preference approach to elicit agricultural landowners' willingness to accept certain levels of financial incentives to install and maintain tree or grass buffers of various widths for 15 years.  A response function, or the probability of buffer installation on agricultural land under different payment levels, is determined with data from a telephone survey conducted in 2000 in the Maryland section of the Chesapeake Bay watershed.  Payments needed to convince landowners to install buffers are hypothesized to vary with agricultural landowners' concerns about lost income, financial incentives provided and the effect of the buffers on real estate development potential. Understanding landowners' willingness to accept financial incentives to install or widen riparian buffers can permit better targeting of outreach and program efforts.  This information can  also be used to assess the costs of obtaining different program goals, such as the 100,000 acre buffer goal of the CREP.

Landowner adoption behavior of other conservation practices has been examined using different approaches.  Ervin and Ervin provide a review of the research between 1950 and 1980 on adoption of soil conservation practices.  This early research led to the introduction in 1985 of the CRP as a national soil conservation program.  The CRP provides rental payments for environmentally sensitive lands that are removed from production.  Several authors have used survey data to examine the determinants of farmer participation in the CRP (Konyar and Osborn; McLean-Meyinsse, Hui and Joseph; Mortensen et al.; Olmstead and McCurdy; Force and Bills; Esseks and Kraft).  Hagan studied determinants of landowner participation in the Maryland Buffer Incentive Program (BIP), finding that gross income from farming and percentage of net income from farming were important differences between participants and nonparticipants.[1] Participants also tended to be younger, have more education, earn less than $1000 from farming, and have fewer years of farm experience.

Contingent valuation has been used to examine the level of incentives needed to induce landowner adoption of conservation practices. Gasson and Potter found that longer-term conservation practices needed higher rental payments. Cooper and Keim used a direct revelation technique, finding that higher incentive levels positively influenced and off-farm work negatively influenced the probability of farmers adopting water quality protection practices. Cooper found that the contingent valuation approach overestimates the minimum incentive payment a farmer would accept to adopt conservation practices when compared to the actual payments that induced participation. Nonparticipants of the BIP said they would need between $730 to $1121 per acre to establish a forest or grass buffer for 10 years (Hagan). Kline, Alig, and Johnson found the mean incentive payment necessary to induce forest owners to forego timber harvest in riparian areas ranged from $38 to $137 per acre annually depending on the owner's objectives (timber versus recreation) for the land. Lohr and Park distinguished the discrete choice to participate in a filter strip program from the continuous choice of how many acres to plant and found that estimated willingness-to-accept payments are not uniform across locations.

Authors also have examined how expectations about land development might affect participation behavior. These studies recognize that landowners in suburbanizing areas must decide whether they would prefer to retain their land without any constraints on conversion – such as those imposed by establishing a riparian buffer and signing a CREP contract – in order to be able to take advantage of an optimal sales opportunity if one should appear. Hansen and Schwartz found that enrollment behavior in the California Williamson Act, a 20-year-contract preferential taxation assessment program, was not independent of the landowner's development expectations. They concluded that landowners' expectations of an exceptionally favorable sale within the 20 year period were overly optimistic. Lynch and Brown found that Maryland landowners in rapidly developing counties where land values are appreciating would not maximize overall net benefits by enrolling in CREP. Mooney and Eisgruber found that installing a riparian forest buffer reduced the market value of Oregon residential parcels with streams.

This analysis differs from previous work in several ways. First, even if all of a farm owner's water bodies have buffers, she is included in the study. A large proportion of water bodies have a narrow strip of trees or grass on their banks. While these narrow strips of grass or trees do provide some environmental benefits, widening them could enhance their effectiveness in achieving water quality and wildlife habitat goals. Consequently, landowners who indicated that all of their water bodies were buffered were asked about their willingness to widen existing buffers by 35 or 100 feet. Second, landowners who had previously installed riparian buffers on their streams were also included in the study. Variables were used to measure the effect of prior knowledge about buffers and previous participation in government programs on owner's willingness to accept payment for the provision of buffers. Third, the offered incentive payments included a range whose high end exceeded the agricultural opportunity cost of the buffered land. The study area within Maryland is undergoing suburban development and agricultural landowners are expected to consider the option value of having unencumbered land available for sale. Fourth, we examine landowners' willingness to accept payments for different buffer widths and different types of vegetation. This is done to investigate how buffer specifications will affect landowner willingness to accept a financial incentive. Fifth, the survey instrument allowed respondents to answer "don't know" to the willingness to accept question. The motivation for these noncommittal responses is assessed through a comparison of the ordered probit and multinomial logit models.

**Modeling Framework**[2]

The analysis is based on random utility theory, which permits discrete choices within a utility-maximizing framework (Hanemann 1984). According to this theory, the landowner's indirect utility function can be represented by $V_i\,(x_i, q)$ where $x_i$ is a vector of measurable personal and land characteristics of landowner $i$ and $q$ indicates the buffer installation choice: $q = 0$ if no buffer is installed (the status quo) and $q = 1$ if the buffer is installed.[33] Utility maximization implies that a landowner chooses not to plant a buffer if it will not increase his indirect utility, $V_i(x_i, 0) > V_i\,(x_i, 1)$, and chooses to install a buffer if doing so causes the greater utility, $V_i\,(x_i, 0) < V_i\,(x_i, 1)$. Benefits from buffer installation can be increased by providing incentives, and installation will occur if the level of monetary incentive is large enough to ensure that the individual's indirect utility from buffer installation is greater than the utility under the status quo. The lower bound for this incentive is $C_i$, which is defined as the value that makes $V_i\,(x_i, 0) = V_i\,(x_i, 1, C_i\,)$. Thus an owner's willingness to plant the buffer can be altered by offering an incentive bid level $A$ that ensures $V_i\,(x_i, 0) < V_i\,(x_i, 1, A\,)$, i.e., that $A > C_i$. If this incentive is offered in a stated preference situation, a landowner should respond that he will accept the payment and install the buffer. Similarly, if the costs remain larger than the benefits, $V_i\,(x_i, 0\,) > V_i\,(x_i, 1, A\,)$, then a landowner should say he is unwilling to accept the incentive payment. The incentive level offered could result in a situation where $V_i\,(x_i,\,0) = V_i\,(x_i,\,1,A)$, and the owner may indicate this indifference by saying she does not know if she will accept the incentive and plant the buffer.

The stated preference question used to elicit landowner preferences for buffer installation is for a specific annual rental amount to be paid for a 15 year period. Since riparian buffers are heterogeneous in nature, landowners will visualize buffers with different characteristics when answering this question. To remove some of this heterogeneity, buffer width and vegetative type are specified in the question. Different combinations of width and type provide a range of responses and a "supply" function for the various characteristics of the buffers. The contingent valuation survey question was preceded with the statement: "The next few questions are about government programs that help farm landowners install new or widen existing buffers. These questions concern alternative program options that might be offered." Landowners who had water bodies without buffers were asked:

> "These programs would help farm landowners install riparian buffers. One option would cover 100% of the costs of a (tree, grass) buffer that averages (35, 100) feet in width from the water's edge. It would provide an annual payment of (40, 90, 140, 190) dollars per acre for 15 years. If available, would you participate in this program?"

Landowners who had buffers for all of their water bodies were asked:

> "These programs would help farm landowners widen existing buffers. One option would cover 100% of the costs of a (tree, grass) buffer that averages (35, 100) feet more than the existing buffer. It would provide an annual payment of (40, 90, 140, 190) dollars per acre for 15 years. If available, would you participate in this program?

The options of tree or grass, 35 or 100 feet, and 40, 90, 140 or 190 dollars were randomly assigned to survey respondents so that roughly equal proportions of responses were obtained for each combination of options.

**Econometric Models**

In the random utility framework, the minimum willingness to accept, $C_i\ (x_i\ )$, is treated as a random variable. While individual landowners are assumed to know their minimum willingness to accept with certainty, these values are revealed only through observed choices or elicited preferences, and determinants of these values often are unobservable, observable with error or measurable only through proxies. Thus, an observer cannot infer $C_i\ (x_i)$ with certainty. Cameron incorporates this uncertainty by specifying a cumulative distribution function $G_C$ for the minimum random willingness to accept value. Then the probability that a landowner will answer "yes" to the contingent valuation question will be $Pr(\ Yes) = G_C\ (A)$. If the expected value of $C_i(x_i\ )$ is a linear function of the value's determinants ( i.e., $\mu = x_i\, b$ where $b$ is a vector of parameters), and the variance of $G_C$ equals $s\ ^2$, one may specify $G$ to be the cumulative distribution function for the standardized variate $z = (C_i\ \text{-}\ x_i\, b)\ /\ s$. Then $Pr(\ Yes) = G\ (?A\ \text{-}\ x_i\, ß)$ where $? = 1/s$ and $ß = b/s$. Estimates of the parameters $?$ and $ß$ can be obtained once the form of $G(z)$ is specified. For example, if $G(z)$ is assumed to be the standard normal cumulative distribution function F$(z)$, one can estimate the parameters $?$ and $ß$ using a probit model, and if $G(z)$ is assumed to be the standard logistic function $(1\ +\ e^{\text{-}\ y})^{\text{-}\ 1}$, one can estimate the parameters using a logit model. Other distributional assumptions are possible (Maddala).

Concerns about respondents' ability to state their true willingness to accept or pay have caused debate about the stated preference method. Respondents may be unable to provide their true preferences because they have had little prior experience with the item in question and thus have difficulty establishing their minimum willingness to accept during a single survey (Cummings et al.). A person's willingness value may be formed or adjusted by the new information provided by the survey itself (Gregory and Slovic). Moreover, if the offered incentive is truly a threshold bid for individual *i*, then "don't know" or "indifferent" could be the valid answer. How the item is defined in the contingent valuation question also may affect elicited responses. If the item is vaguely defined, respondents may be unsure how to respond. The true response may be "yes" if the item has certain characteristics but "no" if it has other characteristics. This answer may translate into "don't know" when the only option is to respond yes, no, or don't know. Svento found that recoding "don't know" answers into the "no" category or into the "yes" category can result in substantially different aggregate benefit measures for a project. Using an ordered probit model did not decrease the variance of the estimates but did permit Svento to tease out the "indifference belt" around the yes and no answers. Thus, the model may be better constructed as $V_i(x_i, 0, d) < V_i(x_i, 1, C)$, where the parameter d measures an interval around an indifference curve where the landowner is uncertain about the utility derived from installing a buffer and consequently uncertain about whether to answer "yes" or "no." Hanemann and Kanninen refer to this specification as a model of "thick" indifference curves and trace it back to Georgescu-Roegen, Luce and Quandt.

Twenty percent of the surveyed landowners responded with "don't know" to the buffer installation question. Rather than ignoring these "don't know" responses, three different econometric models are estimated. We estimate a binary probit model in which the "don't know" responses are treated as missing answers (Yes=1; No =0).[44] Second, we estimate an ordered probit model assuming there is a natural ordering of the discrete choices. Landowners who are willing to install have a high probability, i.e., their true willingness, $z$, will be high. Conversely, landowners who receive less utility or are less willing will have low values for $z$. If landowners were indifferent to installing a buffer, their true willingness should fall between those willing and those unwilling (Yes=2; Don't know=1; No=0). A multinomial logit regression is also estimated to test the ordering assumption and to determine if the characteristics have marginally different impacts on willing and uncertain landowners.

For the ordered case, the stated preference, $y_i$, is used as the dependent variable assuming that

$y_i = 0$ (*unwilling to accept)* if $a_{-1} < z_i < a_0$;

$y_i = 1$ (*don't know*) if $a_0 < z_i < a_1$ ;

$y_i = 2$ (*willing to accept)* if $a_1 < z_i < a_2$;

where $a_{-1} < a_0 < a_1 < a_2$ . The a's are parameters that bound the ranges containing the true willingness value, $z_i$. No significance is assigned to the unit of distance between the stated responses, $y_i$'s: $a_{-1}$ is set = - 4 , $a_2 = +4$, and $a_0$ is anchored at zero. The parameter $a_1$ is estimated as part of the model estimation process. The Prob($y_i = j$) is the probability that $z_i$ is in the $j^{th}$ range, and $z_i$ is assumed to be within the $j^{th}$ range if $a_{j-1} < z_i < a_j$ (j = 0,1, 2).

The probability of recording $j$ as an individual's response is

$Prob(y_{ij} = 1) = F(a_j \& x_i ß \% A?) \& F(a_{j\&1} \& x_i ß \% A?)$, where $\Phi$ is the cumulative density function for the standard normal distribution. As before, $x_i$ is a vector of exogenous characteristics of individual $i$, $A$ is the offered payment level, and $a_1$, $\$$ and ? are coefficients to be estimated. Coefficients can be estimated using the likelihood function: $L \,' \mathop{J}_{i} \mathop{J}_{j} [F(a_j \& x_i ß \% A?) \& F(a_{j\&1} \& x_i ß \% A?)]^{y_{ij}}$.

The multinomial logit model uses the stated preference $y_{ij}$ as the dependent variable assuming that

$y_{ij} = 0$ (*unwilling to accept*) if $z_{i0} > z_{i1}, z_{i2}$ ;

$y_{ij} = 1$ (*don't know*) if $z_{i1} > z_{i0}, z_{i2}$ ;

$y_{ij} = 2$ (*willing to accept*) if $z_{i2} > z_{i0}, z_{i1}$.

The probability that an individual's response was answer *j*, where *j* = 0,1, 2, is $Prob(y_{ij} = j) = \frac{e^{A?_j \& x_i ß_j}}{\sum_j [e^{A?_j \& x_i ß_j}]}$. In this formulation, elicited responses are not bounded but parameter estimates are restricted to the relative measures. We used Limdep 7.1 to compute the regression estimates.

**Data**

A telephone survey of 506 farmland owners was conducted in the spring of 2000 to test the model. A sample of 1032 agricultural landowners, drawn from the Maryland Assessment and Taxation Database, was stratified by four geographic regions of the state. No phone number could be found for 274 individuals. 103 people were ineligible because they do not own a farm or because they have no water bodies on their farm. Thus, there were 655 landowners who were deemed eligible from the initial list. Seventy-four contacted individuals refused to participate. This was 15 percent of the final completed surveys.

Table 1 contains variable definitions. Descriptive statistics for the study variables are reported in Table 2. The average payment offered to the respondents was $112. Willing respondents have a significantly higher average payment of $120 compared to the average payment of $103 for unwilling respondents, consistent with our theoretical model.[55]

Although the differences between mean education and age are small for willing and unwilling respondents, both differences are statistically significant. In contrast, no significant differences are found in the average farm size. The acreage distribution is skewed to the right with an average of 181 and a median of 97.5 acres. Almost eighty percent of the survey respondents had heard about riparian buffers with more yes respondents (87%), fewer no respondents (79%) and even fewer "don't know" respondents (67%) similar to what would be expected from the Cummings et al. hypothesis.

Two dependent variables are defined: WILLDKP (Yes=1; No=0), for which "don't know" responses are treated as missing values in the binary probit, and WILLDKO (Yes=2; Don't know=1; No=0), for the ordered probit and the multinomial estimation.

PAYMENT is the level of per acre annual incentive payment ($40, $90, $140, $190). It corresponds to *A* in the random utility model; thus it is one benefit a landowner receives by planting a buffer. The width of the buffer (SIZE) affects the placing of the buffers on the land and its potential impact on agricultural income. This variable measures both the amount of land taken out of production and the total revenue gained from the buffer incentive payment. Since these represent offsetting effects on the landowner's decision, the sign and significance of the coefficient on SIZE will depend on which effect dominates. The vegetative type of buffer (TREES) may affect an owner's willingness to install a buffer. Planting a tree buffer can increase the conversion costs from farmland to residential, industrial or commercial use if the trees need to be cleared or protected during development. Although residential lots adjoining parcels with trees can have higher sales values, if the landowner expects to convert the land in the future, he is hypothesized to be less willing to plant a tree buffer. Landowners who already have buffered their water bodies are assumed to have lower transaction costs to widen these existing buffers, thus are hypothesized to be more willing to accept a financial incentive. These owners are distinguished by the variable INSTALL .

Direct measures of the lost agricultural income associated with installation of the buffer were unavailable on an individual farm basis. This opportunity cost is proxied using the farm acreage (ACRES) and the width of the buffer (SIZE). Landowners with larger farms are expected to be more willing to install riparian buffers since doing so will affect the scale of the farm operations less.

Four additional variables address the issue of whether the landowner is also the farmer. Respondents who do not make farming decisions or whose land is not used for farming are indicated by the binary variable NOTFARM. The variable OWNDEC indicates whether the respondent makes farming decisions for the farm property on a unilateral basis. A landowner who makes farm production decisions solely, without consulting a partner, tenant or other family member, is expected to be more willing to install a buffer for two reasons. One, she could commit to buffer installation on the phone if she is the sole decision-maker for the farm property. Two, this person may have a better sense of the lost agricultural income relative to the offered incentive payment. The variable LOWINC measures whether 1-24 percent of the owner's income comes from farming. Landowners who gain most of their income from off-farm activities should be less concerned about the loss of agricultural income from installing buffers and are expected to be more willing to install a buffer. If these landowners lease out their land, however, they may be concerned about their tenant's willingness to continue renting the property if fewer acres were available to farm. NOINC equals one if none of the landowner's income comes from farming. Such owners could be land investors who are concerned with increases in land value, retired farmers, and/or people who receive non-consumptive value from owning and living on a farm. NOINC tests whether landowners who hold farmland for non-farming purposes are more or less willing to accept a payment for installing buffers.

As in many indirect utility models, EDUC and AGE are included to account for differences in utility functions across respondents. In addition, higher education should decrease the transaction costs of learning about buffers and incentive programs. Therefore, respondents with higher educations are expected to be more willing to install buffers. If an older farmer wants to sell his farm to finance retirement, then he would not wish to encumber the land with a 15-year buffer contract. Younger farmers may see a longer time horizon to benefit from the buffers, especially if they agree to plant a forest buffer. KEEPFARM equals 1 if the landowner intends to keep farming for more than the15-year contract period. Farmers who plan to keep farming past the contract period would not care about potential development options during that period. AGE is expected to decrease willingness, and KEEPFARM is expected to increase willingness to install a buffer.

We also include binary variables for regions of Maryland. Regions represented by binary variables include the Upper Eastern Shore (UESHORE), Lower Eastern Shore (LESHORE), and Southern Maryland (SOUTH). Counties of Central Maryland and Western Maryland are represented by zero values for the three regional binary variables. Regions have different soil types and micro climates that affect agricultural net revenues. In addition to being proxies for agricultural income, the region variables also serve as proxies for development potential (timing) and value. Central Maryland is experiencing relatively more development and the value of converting farms to non-agricultural uses is increasing in this region. Southern Maryland has also been experiencing rapid growth. Since agricultural productivity and development prospects can have countervailing effects on willingness to install buffers, estimated coefficients for the region variables could be positive or negative.

Transaction costs are expected to directly affect landowners' willingness to install riparian buffers. Besides the influence of OWNDEC and EDUC, three additional variables (GOVTPROG, EXTED, KNOW) are included to represent these costs. Landowners who have participated in a government program within the past five years, such as Environmental Quality Improvement Program or Conservation Reserve Program, are expected to be more willing to participate in a buffer installation program. These landowners have already visited the county offices and thus have lower learning costs and other transaction costs as some of their information will be on file. Landowners not already enrolled in government programs may be ineligible to participate or may have an aversion to participating in government programs. Landowners who recalled receiving educational material about riparian buffers from Maryland Cooperative Extension (EXTED) also are expected to be more willing to agree to install a buffer. These educational materials, provided as a separate part of the survey project, described riparian buffers, their benefits and costs, and the government programs available to support their installation. Having such material delivered to one's home should decrease the transaction costs of learning about buffer installation. KNOW equals one if the respondent has heard of riparian buffers. Costs of learning about buffers will be lower for these owners. Uninformed landowners may not have the necessary information to determine their minimum willingness to participate in a buffer program. Uninformed owners are expected to be less likely to agree to install a buffer. They are more likely to say "don't know".

Some respondents would not give the percent of income from farming, others did not specify the number of acres owned, and others did not indicate how long they planned to keep farming. In order to make use of these observations, binary variables were created that were equal to one if the data was missing (zero otherwise) and then the relevant variables' missing values were changed to zero.

These binary variables were included in the models to determine if the missing value transformations had a significant effect on the model parameter estimates. The missing value variables included in the models are DLOWINC, DKEEP, DEXTED, and DGPROG.

**Results**

Coefficient estimates for a probit model in which "don't know" responses are treated as missing (WILLKNP is the dependent variable) and an ordered probit model in which "don't know" responses are treated as a separate category (WILLKNO is the dependent variable) are reported in Table 3. This table also provides the estimated value of $a_1$, the interval value for the ordered probit. Including the independent variables listed in Table 1 significantly increased the likelihood values in both models over a model in which all slopes are set equal to zero.[6] Overall, the two models correctly predict about two-thirds of the "yes" and "no" responses, with the ordered probit doing slightly better on the "yes" responses and slightly worse on the "no" responses.[7]

A multinomial logit model also was estimated, in part to assess the effect of forgoing the ordering assumption of the ordered probit model. This model, presented in Table 4, is significantly also better than a model which sets all slopes equal to zero ($?^2 = 105.06$, $d.f. = 42$), and shows significant differences in parameter estimates between the three types of responses. This model correctly predicts 78 percent of the "no" responses and 50 percent of the "yes" responses, but only 26 percent of the "don't know" responses. In comparison, the ordered probit predicts 79% of the "no" responses, 48% of the "yes" responses, and none of the "don't know" responses. Thus imposing the "thick indifference curve" hypothesis under which the "don't know" response is between "yes" and "no" does not improve the model's predictive capabilities. People may not know how willing they are for a variety of reasons:

for some it could be they are indifferent to the offered incentive, for others they may not be able to evaluate the offered contract due to lack of knowledge. While the models explain “yes” and “no” answers, they do not perform well for the “don’t know” responses.

As hypothesized, increases in the incentive payment, *A*, make landowners more likely to agree to install a buffer. All three models produce significant coefficient estimates for PAYMENT. These coefficients range from 0.0032 for the ordered probit model to a comparable 0.0036 for both the binary probit and the multinomial logit model.[88] The marginal values measured at the $110 mean bid value are close in magnitude: 0.00138 for the binary probit model, 0.0011 for the “yes” response in the ordered probit model and 0.00105 for the “yes” response in the multinomial logit model (Tables 5-6). Thus, responses to changes in incentive payments are quite robust across models for middle level bid values. The estimates indicate that the probability that a landowner will respond “yes” to the buffer installation question will increase by 1.0% to 1.4% , given a marginal increase in the $110 payment.

The models predict the same willingness to accept payment for installing a riparian buffer, regardless of the specified width of buffer (100 vs. 35 feet) or type of vegetation (trees vs. grass). We have experimented with a “fully saturated” hedonic utility formulation (Hanemann and Kanninen, pp. 355-58) but found that separate questions for each buffer type did not predict different probabilities of “yes” responses for buffers with different characteristics. Accordingly, we include TREES and SIZE simply as intercept shifters. As Tables 3 and 4 show, neither variable has a significant effect on the willingness to adopt.[99] For illustrative purposes, we graphed the average probability for grass and tree buffers against the four bid levels in Figure 1.[1010] From this figure, we hypothesize that the respondents’ apparent reluctance to adopt grass buffers at a low bid level may explain this unexpected result.

The estimated model does not support the hypothesis that predictions are uniformly higher or lower for respondents who stated that all of their water bodies have buffers; the coefficients on INSTALL are insignificant. This suggests that adoption behavior is similar even for those who already have adopted buffers on all of their water bodies. Perhaps some of these landowners have already planted sufficiently wide buffers and do not want any more added while other landowners are agreeing because they have found buffers beneficial.

Of our proxies for lost agricultural income, only the estimated coefficient on low farm income (1-25% from farming) is significant. Landowners with a low percentage of family income from farming are more willing to install a buffer compared to landowners who obtain more than 25% of their income from farming. This suggests that the loss of the agricultural income or flexibility may be more important to full-time farmers and the offered incentive levels may not be high enough to compensate them for this loss. Landowners with no income from farming behave similarly to the full-time farmers. This result could occur because these owners are concerned about their tenants' willingness to keep renting the land, because they have significantly higher transaction costs when enrolling in a buffer program, or because they want to preserve their development option.

Development value proxies behaved contrary to expectations. All three models indicate that farmers who plan to keep the farm for 15 years or more are less likely to install a buffer than those who do not have such plans. Respondents who plan to keep the farm for 15 or more years were 14.6 to 15.6 percent less likely to say "yes" than those who did not plan to keep the farm for this time period. Farmers who do not plan to keep farming for the entire 15 year contract apparently do not find the loss of development options critical.

Age also negatively affects a landowner's willingness to install a buffer Older landowners may want to sell the property to finance their retirement, want their children to have flexibility, or may think they will not benefit for the full length of the contract. An additional year of age decreases the probability of agreeing to install a buffer by 4.8 to 6.7 percent. Figure 2 shows the average probability of agreeing to different bid levels by 4 categories of age. This figure clearly illustrates that older landowners are less willing to install a buffer at every bid level. Younger farmers (27-50 years old) had an average probability of a "yes" response of more than 60%, while the oldest age group (76 +) had a probability of less than 40%. Individuals with more years of education are more likely to respond "yes" in the binary and multinomial models. A marginal increase in one's education increases the likelihood of a "yes" response by 1.7 to 2.1 percent.

The regional variables included show that farmers in Southern Maryland are less likely to agree to install a buffer, while farmers in the Lower Eastern Shore are more likely. For Southern Maryland, the loss of development potential appears to outweigh the level of the incentive payment. In the Lower Shore, where development pressure is lower, the incentive payment may be sufficient compensation for the lost agricultural income. Lower Shore lands are also more likely to have many water bodies, thus a large number of eligible acres. Figure 3 demonstrates the dominance of the Lower Eastern Shore landowners' average probability of enrolling at each level of the incentive payment relative to landowners in the Central and Southern regions.

None of the dummy variables included to represent missing data are significant. Thus adjusting these variables to allow the inclusion of these observations does not affect the estimated coefficients.

**Concluding Remarks**

All the models indicate that the likelihood of agreeing to install a buffer increases with the level of the incentive payment, with a low percentage of income from farming, and for landowners located in the Lower Shore counties. They also indicate that the likelihood decreases for those landowners planning to farm for more than 15 years, older farmers, and those in Southern Maryland. The type of vegetation, the size of the buffer, the number of acres owned, and Upper Eastern Shore location are not important in determining the probability of agreeing to plant a buffer. Estimating models that include the "don't know" responses provides some additional information, although less than we anticipated. Landowners who make farm decisions with others, with lower education, who have received extension materials, and who are located in the Upper Eastern Shore and South Maryland, are all more likely to say "no" compared to "don't know." The multinomial logit under which the "don't know" responses are treated as a separate, but not an ordered, category, has the most similarity to the binary probit where the "don't know" respondents are treated as missing. We find we cannot conclude that "don't know" is a middle category, i.e., that the ordering assumption employed in the ordered probit model may not be accurate.

The most consistent and important result is that the probability of a "yes" response could be increased 1.1% to 1.4% by a marginal increase in the payment rate. Clearly, this is the easiest method of enrolling more land. However, there exists a prevailing philosophy that these programs should only compensate landowners for the lost agricultural income and not for development value. Maryland crop budget estimates for 1999 indicate that the net revenue (excluding land rental) for corn silage is $119/acre and for a wheat/soybean rotation $124/acre. The Maryland 2000 average rental rate for farmland was $54.20 per acre, ranging from $26 to $86 per acre. Thus for most areas of the state, the

$90, $140 and $190 incentive payments would have been more than adequate to compensate landowners for lost net agricultural income or lost land rental income. This suggests that other opportunity or transaction costs are affecting a landowner's willingness to accept payment for installing a riparian buffer.

Full-time farmers appear to be the least likely to agree to take land out of production and to install buffers. Before increasing payments to these farmers, scientists need to examine whether the long-run environment benefits of buffers on the water bodies of full-time farmers would decrease the potential nutrient run-off and leaching more than buffers on a part-time farmers' riparian areas. If the potential water quality improvement is sufficiently high, enrolling these farmers at a higher annual rental payment may be cost effective. Other incentive mechanisms may also be needed to ensure farmer participation.[11]

This result does not bode well for proposed provisions for the 2002 Farm Bill. Several bills suggest a shift of funds to environmental programs to provide an income safety net for farmers while achieving environmental goals. If full-time farmers are less likely to join enhanced conservation programs, then this shift of funds will not be an effective way to achieve both goals.

Targeting Lower Eastern Shore landowners should elicit the most enrollment for Maryland's program. This area houses the majority of the chicken houses in the state and is being targeted by other programs to manage their nutrient loadings more effectively. Thus, installation of riparian buffers is very timely. Attempts have been made to encourage more enrollment in the Southern counties, but these survey results suggest there is less likelihood of high enrollment levels in this part of the state. Broadly speaking, targeting areas with high levels of agricultural nonpoint nutrient pollution and with low development pressure may result in the biggest impact per dollar spent. Furthermore, targeting areas with many water bodies would result in more participation as the transaction costs of enrollment would be lower on a per acre of eligible land basis.

Farmers with pre-existing buffers behaved similarly to those who did not have buffers. This suggests farmers will widen existing buffers under the same conditions and incentives as those needed to motivate first time installation. While the size of the buffer and the vegetative type were not significant in explaining landowner survey responses, respondents were not given a continuous choice. Instead they were offered a complete contract and asked if they would be willing to accept this or not. Riparian buffer workshops indicate that farmers are concerned about the width requirements and that trees are thought to have more drawbacks, including shading the field, increasing deer presence, and restricting the movement of farm equipment (Eastern Shore Tributary Teams, Lynch and Brown). Sign-ups for the CREP program in Maryland indicate a preference for wider buffers and grass over tree buffers, even though the trees pay a higher incentive payment. As one can see in Figure 1, the tree buffer supply curve does lie below the grass buffer curve for higher incentive payments. On the other hand, from December 2000 to December 2001, tree buffer enrollment almost doubled from 5,945 to 11,610 acres while grass buffers increased only 54% from 10,585 to 16,252 acres, suggesting the perception of tree buffers may be changing.

The analysis demonstrates that non-agricultural opportunity costs may be important in the buffer program participation decision. If environmentally sensitive but high-value land, such as in the Chesapeake Bay watershed, is targeted for buffer incentive programs, incentives that more than compensate for agricultural opportunity costs may be necessary to encourage participation. Whether the environmental benefits of enrolling high-valued agricultural lands or incorporating non-agricultural opportunity costs into the rental payment scheme outweigh the additional costs remains to be determined.

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**Footnotes**

1. The Maryland Buffer Incentive Program pays a one-time $300 per acre grant to landowners who establish riparian forest buffers on their property.

2. This explanation of the random utility specification and associated econometric model parallels Hanemann and Kanninen, pp. 307-310.

3. Because an individual's income does not constrain willingness to accept as it does willingness to pay, we do not include it explicitly in the indirect utility function but do include it in the vector of individual characteristics, $x_i$.

4. We also estimated a binary probit model in which the "don't know" responses were assumed to be "no" answers. This model did not perform as well as the binary probit with missing responses; thus for brevity's sake, we do not report it here.

5. Statistical significance of observed differences in mean values of continuous variables between respondents who answered "yes" and who answered "no"are determined using a t-test and a 5 percent level of significance. Statistical significance of differences in the frequencies of "yes" and "no" responses relative to the given proportions of respondents in the discrete variables are assessed using a contingency table, a chi square statistic, and a 5 percent level of significance.

6. $?^2 = 53.4$ and $d.f. = 22$ for the binary probit model, and $?^2 = 53.02$ and $d.f. = 22$ for the ordered probit model.

7. When probability estimates above 0.5 are assigned a "yes" value and estimates below 0.5 are assigned a "no", the binary probit model correctly predicts 155 of 189 no responses and 59 of 129 yes responses. When the threshold level is set equal to the observed frequency of yes responses, as suggested by Khanna and described by Greene (p. 833), the binary probit model correctly predicts 124 of the 189 no responses and 88 of the 129 yes responses. The ordered probit correctly predicts 150 of the 189 no responses and 62 of the 129 yes responses.

8. Following Amemiya, the logit estimate is multiplied by 0.625 to make it comparable to the probit estimates (Maddala, p. 23).

9. In reality, fewer acres of trees have been enrolled in the Maryland CREP program even though the incentives and cost-share rate are higher.

10. The predicted probabilities in all of the figures in the paper are computed using the binary probit model.

11. For example, forest landowners indicate a preference for property tax forgiveness rather than incentive payments to apply poultry litter on their land (Lynch and Tjaden 2001).

**Table 1: Definition of Variables**

| **Variable** | Description |
|---|---|
| ***Dependent*** | |
| **WILLDKP** | Willing to Install Buffer: probit (=1 if yes, 0 if no or don't know ) |
| **WILLDKO** | Willing to Install Buffer: others (=2 if yes, 1 if don't know, 0 if no) |
| ***Independent*** | |
| **PAYMENT** | Prices (40, 90, 140, 190) randomly assigned ($) |
| **TREES** | Tree used as buffer (=1 if yes, 0 if grass) |
| **SIZE** | Buffer size is 100 (=1 if yes, 0 if 35) |
| **INSTALL** | Install buffer (=1 if install, 0 if widen) |
| **OWNDEC** | Decision production is made by owner alone (=1 if yes) |
| **ACRES** | Farm acres owned |
| **NOINC** | Percent of household income from farming is zero (= 1 if yes) |
| **LOWINC** | Percent of household income from farming is between 1-25% (=1 if yes) |
| **KEEPFARM** | Farmers plan to keep farming more than 15 years (=1 if yes) |
| **NOTFARM** | Respondent's land is not used for farming and no farm decision made |
| **KNOW** | Respondents have heard about riparian buffers (=1 if yes) |
| **GOVTPROG** | Participated in a government program in last 5 years (=1 if yes) |
| **EXTED** | Educational materials about buffers received (=1 if yes) |
| **EDUC** | Years in School |
| **AGE** | Age of respondents in years |
| **UESHORE** | Parcel in Talbot, Caroline, Queen Anne or Kent counties (=1 if yes) |
| **LESHORE** | Parcel in Somerset, Worcester, Wicomico or Dorchester counties (=1 if yes) |
| **SOUTH** | Parcel in Calvert, Charles, or St. Mary's counties (=1 if yes) |
| **DGPROG** | Dummy equals 1 if GOVTPROG is missing |
| **DKEEP** | Dummy equals 1 if KEEPFARM is missing |
| **DEXTED** | Dummy equals 1 if EXTED is missing |
| **DLOWINC** | Dummy equals 1 if LOWINC or NOINC is missing |

**Table 2. Descriptive Statistics**

| Variables | All Respondents (N=453) | | Yes responses (N=129) | | No responses (N=189) | | Don't know responses (N=72) | |
|---|---|---|---|---|---|---|---|---|
| | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. |
| **PAYMENT** | $112.296 | 55.551 | $119.845 | 53.224 | $103.492 | 55.930 | $117.083 | 53.691 |
| **SIZE** | 0.499 | 0.501 | 0.512 | 0.502 | 0.524 | 0.501 | 0.431 | 0.499 |
| **TREES** | 0.512 | 0.500 | 0.481 | 0.502 | 0.524 | 0.501 | 0.514 | 0.503 |
| **INSTALL** | 0.317 | 0.466 | 0.341 | 0.476 | 0.328 | 0.471 | 0.500 | 0.707 |
| ***Agricultural Variables*** | | | | | | | | |
| **OWNDEC** | 0.459 | 0.499 | 0.465 | 0.501 | 0.529 | 0.500 | 0.389 | 0.491 |
| **ACRES** | 181.130 | 276.495 | 226.240 | 372.766 | 176.751 | 253.389 | 112.083 | 183.453 |
| **NOINC** | 0.249 | 0.433 | 0.186 | 0.391 | 0.302 | 0.460 | 0.277 | 0.451 |
| **LOWINC** | 0.366 | 0.482 | 0.481 | 0.502 | 0.302 | 0.460 | 0.292 | 0.458 |
| **KEEPFARM** | 0.351 | 0.478 | 0.326 | 0.470 | 0.397 | 0.491 | 0.222 | 0.419 |
| **NOTFARM** | 0.294 | 0.456 | 0.271 | 0.446 | 0.243 | 0.430 | 0.444 | 0.500 |
| ***Transaction Costs*** | | | | | | | | |
| **KNOW** | 0.786 | 0.411 | 0.868 | 0.340 | 0.788 | 0.410 | 0.667 | 0.475 |
| **GOVTPROG** | 0.338 | 0.473 | 0.426 | 0.496 | 0.302 | 0.460 | 0.194 | 0.399 |
| **EXTED** | 0.364 | 0.482 | 0.488 | 0.502 | 0.365 | 0.483 | 0.236 | 0.428 |
| ***Demographics*** | | | | | | | | |
| **EDUC** | 13.130 | 3.546 | 14.140 | 2.861 | 13.148 | 3.594 | 11.403 | 4.740 |
| **AGE** | 61.360 | 13.864 | 58.659 | 13.745 | 61.730 | 13.645 | 61.386 | 17.534 |
| **UESHORE** | 0.223 | 0.417 | 0.248 | 0.434 | 0.249 | 0.433 | 0.153 | 0.362 |
| **LESHORE** | 0.132 | 0.339 | 0.186 | 0.391 | 0.079 | 0.271 | 0.278 | 0.451 |
| **SOUTH** | 0.073 | 0.260 | 0.039 | 0.194 | 0.111 | 0.315 | 0.083 | 0.278 |
| ***Binary Variables for Missing Values*** | | | | | | | | |
| **DLOWINC** | 0.084 | 0.278 | 0.031 | 0.174 | 0.069 | 0.254 | 0.236 | 0.428 |
| **DGPROG** | 0.026 | 0.161 | 0.031 | 0.174 | 0.021 | 0.144 | 0.042 | 0.201 |
| **DKEEP** | 0.066 | 0.249 | 0.062 | 0.242 | 0.085 | 0.279 | 0.069 | 0.256 |
| **DEXTED** | 0.196 | 0.398 | 0.147 | 0.356 | 0.217 | 0.413 | 0.250 | 0.436 |

**Table 3. Coefficient Estimates for Binary and Ordered Probit Models Explaining Willingness to Accept Payment to Install or Widen a Riparian Buffer[a]**

| | Binary Probit (N=318) | | Ordered Probit (N=396) | |
|---|---|---|---|---|
| **Variables** | Coefficient | Std.Err. | Coefficient | Std.Err. |
| **Constant** | -0.826 | 0.634 | -0.056 | 0.502 |
| **PAYMENT** | 0.004 ** | 0.001 | 0.003 ** | 0.001 |
| **SIZE** | 0.111 | 0.157 | 0.075 | 0.126 |
| **TREES** | -0.065 | 0.155 | -0.039 | 0.124 |
| **INSTALL** | 0.059 | 0.165 | 0.025 | 0.134 |
| ***Agricultural Variables*** | | | | |
| **OWNDEC** | -0.117 | 0.180 | -0.192 | 0.148 |
| **ACRES** | 0.000 | 0.000 | 0.000 | 0.000 |
| **NOINC** | -0.130 | 0.234 | -0.097 | 0.188 |
| **LOWINC** | 0.478 ** | 0.204 | 0.425 ** | 0.164 |
| **KEEPFARM** | -0.381 ** | 0.195 | -0.29* | 0.304 |
| **NOTFARM** | -0.225 | 0.239 | -0.194 | 0.194 |
| ***Transaction Costs*** | | | | |
| **KNOW** | 0.173 | 0.232 | 0.093 | 0.185 |
| **GOVTPROG** | 0.205 | 0.173 | 0.211 | 0.139 |
| **EXTED** | 0.105 | 0.183 | 0.016 | 0.149 |
| ***Demographics*** | | | | |
| **EDUC** | 0.055 ** | 0.027 | 0.025 | 0.021 |
| **AGE** | -0.013 ** | 0.006 | -0.010 ** | 0.005 |
| **UESHORE** | -0.058 | 0.192 | -0.104 | 0.156 |
| **LESHORE** | 0.534 ** | 0.251 | 0.432 ** | 0.197 |
| **SOUTH** | -0.59* | 0.327 | -0.578 ** | 0.261 |
| ***Binary Variables for Missing Values*** | | | | |
| **DLOWINC** | 0.039 | 0.396 | 0.094 | 0.300 |
| **DGPROG** | 0.227 | 0.531 | 0.264 | 0.501 |
| **DKEEP** | -0.251 | 0.331 | -0.253 | 0.279 |
| **DEXTED** | -0.214 | 0.220 | -0.212 | 0.175 |
| **Mu** | | | 0.558 ** | 0.059 |

[a] One and two asterisks indicate that, on the basis of an asymptotic t-test, we reject the null-hypothesis that the coefficient is zero using the 0.10 and 0.05 criterion, respectively.

**Table 4. Estimated Coefficients for the Multinomial Logit Model Explaining Willingness to Accept Payment to Install or Widen a Riparian Buffer**[a]

| | Response | | | |
|---|---|---|---|---|
| | Don't know | | Yes | |
| **Variables** | Coeff. | Std.Err. | Coeff. | Std.Err. |
| **PAYMENT** | 0.004 | 0.003 | 0.006 ** | 0.002 |
| **SIZE** | 0.069 | 0.293 | 0.141 | 0.249 |
| **TREES** | 0.014 | 0.298 | -0.134 | 0.251 |
| **INSTALL** | -0.357 | 0.328 | 0.080 | 0.266 |
| ***Agricultural Variables*** | | | | |
| **OWNDEC** | -0.820 ** | 0.347 | -0.347 | 0.283 |
| **ACRES** | 0.000 | 0.001 | 0.000 | 0.000 |
| **NOINC** | -0.249 | 0.431 | -0.218 | 0.372 |
| **LOWINC** | 0.253 | 0.397 | 0.856 ** | 0.326 |
| **KEEPFARM** | -0.020 | 0.371 | -0.729 ** | 0.304 |
| **NOTFARM** | 0.138 | 0.429 | -0.505 | 0.377 |
| ***Transaction Costs*** | | | | |
| **KNOW** | 0.089 | 0.387 | 0.167 | 0.362 |
| **GOVTPROG** | 0.470 | 0.329 | 0.318 | 0.278 |
| **EXTED** | -0.851 ** | 0.381 | 0.125 | 0.300 |
| ***Demographics*** | | | | |
| **EDUC** | -0.083 ** | 0.038 | 0.06* | 0.034 |
| **AGE** | 0.008 | 0.008 | -0.029 ** | 0.008 |
| **UESHORE** | -0.786 ** | 0.399 | -0.149 | 0.306 |
| **LESHORE** | -0.153 | 0.519 | 0.812 ** | 0.404 |
| **SOUTH** | -1.28* | 0.682 | -1.01* | 0.561 |
| ***Binary Variables for Missing Values*** | | | | |
| **DLOWINC** | 0.565 | 0.552 | -0.143 | 0.662 |
| **DGPROG** | 0.972 | 0.881 | 0.318 | 0.829 |
| **DKEEP** | -0.816 | 0.841 | -0.383 | 0.532 |
| **DEXTED** | -0.416 | 0.386 | -0.383 | 0.361 |

[a] One and two asterisks indicate that, on the basis of an asymptotic t-test, we reject the null-hypothesis that the coefficient is zero using the 0.10 and 0.05 criterion, respectively.

**Table 5. Marginal Values for the Binary and Ordered Probit**

| | Binary | Ordered Probit Response | | |
|---|---|---|---|---|
| **Variables** | | No | Don't know | Yes |
| **Constant** | -0.317 | 0.022 | -0.003 | -0.020 |
| **PAYMENT** | 0.001 | -0.001 | 0.000 | 0.001 |
| **SIZE** | 0.043 | -0.030 | 0.003 | 0.026 |
| **TREES** | -0.025 | 0.016 | -0.002 | -0.014 |
| **INSTALL** | 0.023 | -0.010 | 0.001 | 0.009 |
| ***Agricultural Variables*** | | | | |
| **OWNDEC** | -0.045 | 0.076 | -0.009 | -0.068 |
| **ACRES** | 0.000 | 0.000 | 0.000 | 0.000 |
| **NOINC** | -0.050 | 0.039 | -0.004 | -0.034 |
| **LOWINC** | 0.183 | -0.169 | 0.019 | 0.150 |
| **KEEPFARM** | -0.146 | 0.114 | -0.013 | -0.102 |
| **NOTFARM** | -0.086 | 0.077 | -0.009 | -0.069 |
| ***Transaction Costs*** | | | | |
| **KNOW** | 0.066 | -0.037 | 0.004 | 0.033 |
| **GOVTPROG** | 0.079 | -0.084 | 0.009 | 0.075 |
| **EXTED** | 0.040 | -0.007 | 0.001 | 0.006 |
| ***Demographics*** | | | | |
| **EDUC** | 0.021 | -0.010 | 0.001 | 0.009 |
| **AGE** | -0.005 | 0.004 | 0.000 | -0.003 |
| **UESHORE** | -0.022 | 0.041 | -0.005 | -0.037 |
| **LESHORE** | 0.205 | -0.172 | 0.019 | 0.153 |
| **SOUTH** | -0.227 | 0.230 | -0.026 | -0.204 |
| ***Binary Variables for Missing Values*** | | | | |
| **DLOWINC** | 0.015 | -0.038 | 0.004 | 0.033 |
| **DGPROG** | 0.087 | -0.105 | 0.012 | 0.094 |
| **DKEEP** | -0.096 | 0.101 | -0.011 | -0.090 |
| **DEXTED** | -0.082 | 0.085 | -0.009 | -0.075 |

**Table 6. Marginal Values for the Multinomial Logit**

| | Multinomial Logit Response | | |
|---|---|---|---|
| **Variables** | No | Don't know | Yes |
| **PAYMENT** | -0.001 | 0.000 | 0.001 |
| **SIZE** | -0.029 | 0.002 | 0.027 |
| **TREES** | 0.021 | 0.009 | -0.030 |
| **INSTALL** | 0.018 | -0.055 | 0.036 |
| ***Agricultural Variables*** | | | |
| **OWNDEC** | 0.128 | -0.097 | -0.031 |
| **ACRES** | 0.000 | 0.000 | 0.000 |
| **NOINC** | 0.057 | -0.023 | -0.034 |
| **LOWINC** | -0.161 | -0.010 | 0.171 |
| **KEEPFARM** | 0.120 | 0.036 | -0.156 |
| **NOTFARM** | 0.070 | 0.047 | -0.117 |
| ***Transaction Costs*** | | | |
| **KNOW** | -0.035 | 0.004 | 0.031 |
| **GOVTPROG** | -0.093 | 0.049 | 0.043 |
| **EXTED** | 0.054 | -0.127 | 0.073 |
| ***Demographics*** | | | |
| **EDUC** | -0.002 | -0.015 | 0.017 |
| **AGE** | 0.004 | 0.003 | -0.007 |
| **UESHORE** | 0.093 | -0.103 | 0.010 |
| **LESHORE** | -0.118 | -0.065 | 0.184 |
| **SOUTH** | 0.275 | -0.126 | -0.149 |
| ***Binary Variables for Missing Values*** | | | |
| **DLOWINC** | -0.026 | 0.087 | -0.061 |
| **DGPROG** | -0.137 | 0.120 | 0.017 |
| **DKEEP** | 0.133 | -0.095 | -0.039 |
| **DEXTED** | 0.099 | -0.038 | -0.060 |

**Figure 1. Predicted Responses by Bid level**