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# HYPERBOLIC DISCOUNTING OF PUBLIC GOODS

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## ABSTRACT

This article examines revealed rates of time preference for public goods, using environmental quality as the case study. A nationally representative panel-based sample of 2,914 respondents considered a series of 5 conjoint policy choices, yielding 14,570 decisions. Both the conditional fixed effect logit estimates of the random utility model and mixed logit estimates implied that the rate of time preference is very high for immediate improvements and drops off substantially thereafter, which is inconsistent with exponential discounting but consistent with hyperbolic discounting. The implied marginal rate of time preference declines and then rises. Estimates of the quasi-hyperbolic discounting parameter range from 0.48 to 0.61. People who are older are especially likely to have a high disutility from delays in improving water quality.

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## 1. Introduction

The rate of discount plays a central role in evaluating many policy decisions. As with other long-term choices, such as those for personal savings, policy choices today may have profound long-term consequences. At the extreme, current policies may have ramifications that will be manifested at the end of this century, as in the case of global climate change. Although there is a substantial literature addressing the normative issue of the appropriate rate of discount for policy evaluation, the behavioral properties of individual rates of time preference are also consequential because these preferences will provide the basis for the public's political support.[1]

Unfortunately, individual rates of time preference may not always exhibit properties associated with the usual intertemporal economic models of rational choice. The conventional exponential discounting model weights utility payoffs in year T by a discount factor $\Phi(T)$, where

$$\Phi(T) = \delta^{T} \tag{1}$$

and $\delta = 1/(1+r)$, where r is the rate of interest. A series of experimental studies that typically have employed student subjects have documented a variety of inconsistencies of the discount rates revealed by individual behavior with the exponential discounting model.[2]

The most widely used alternative to exponential discounting is hyperbolic discounting. Hyperbolic discounting rates place an especially large weight on immediate

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[1] Weitzman (2001) presents a recent contribution to the normative literature on discounting and also presents interesting survey evidence of economic experts regarding the appropriate rate of discount.

[2] For a superb review of this literature, see Frederick, Loewenstein, and O'Donaghue (2002). Recent contributions to the experimental literature include that by Benhabib, Bisin, and Schotter (2004) and Cameron and Gerdes (2005).

payoffs as compared to deferred payoffs, inducing patterns of time inconsistency.[3] The widely used quasi-hyperbolic discounting approach employed by Laibson (1997) is useful because of its analytic simplicity and clear-cut contrast with the exponential model. The quasi-hyperbolic formulation for discrete time periods yields discount factors $\Phi(T)$ given by $\{1, \lambda\delta, \lambda\delta^2, \lambda\delta^3, ...\}$, where $0 < \lambda < 1$, and $\delta < 1$.[4] The discount factor terms involving $\delta$ are all multiplied by a parameter $\lambda$ except in the initial period. The discussion here and below will be in discrete time rather than continuous time because our survey data consists of valuations with respect to four discrete periods of delay.

The approach here will depart from the literature in several ways. First, we estimate rates of time preference based on a series of policy choices administered in a survey context. Rather than using an experimental structure with a small convenience sample, we use a survey methodology drawing on a large nationally representative sample. From this sample we estimate average and marginal rates of time preference and ascertain how these change based on individual characteristics.

Second, we estimate rates of time preference using a random utility model, employing both conditional fixed effect logit and mixed logit approaches. Our estimates of the role of discounting consequently will be based on a theoretically appropriate model of discounted utility rather than money.

Third, our study will test whether the hyperbolic discounting phenomenon is due to the shape of intertemporal preferences or whether uncertainty regarding future payoffs is generating the preference for immediate rewards. Put differently, we will estimate the

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[3] A more flexible formulation of the hyperbolic model proposed by Loewenstein and Prelec (1992), takes the form $\Phi(T) = (1 + \rho T)^{-\omega/\rho}$, where the parameters $\omega$, $\rho > 0$. Concerns with time inconsistency date back to Strotz (1956).

[4] For applications and discussion, see also Angeletos et al. (2001).

quasi-hyperbolic discounting parameter $\lambda$ and explore whether this term is due to the nature of time preference or a perceived probability that deferred payoffs are less likely to be received than immediate payoffs.

Fourth, the experimental structure will pertain to a publicly provided good, environmental quality, rather than a financial outcome. The potential importance of hyperbolic discounting for environmental decisions has been recognized by Cropper and Laibson (1999), and there have been numerous studies of discounting for regulatory outcomes such as lives saved at different times in the future.[5] However, no study to date has presented experimental or empirical evidence pertaining to hyperbolic discounting for publicly provided goods.

Section 2 introduces the policy choice task, which requires that respondents pick the most highly valued alternative from among these policies defined on three dimensions, including improvements in water quality at different times. Section 2 also describes the sample and the validity tests of the methodology. This formulation permits the estimation of a logit model with fixed and random parameters across respondents in Section 3. These models yield information on rates of time preference as a function of different periods of delay. The mixed logit models permit individual heterogeneity in the parameter values and relax two key assumptions of the conditional logit model. We demonstrate the robustness of the results by showing that the two models yield very similar results. Section 4 explores the quasi-hyperbolic discounting parameter $\lambda$ and the issue of whether uncertainty regarding future payoffs is the influence generating the apparent hyperbolic discounting pattern. In Section 5 we explore the variables that

[5] Studies in this vein using surveys or experiments include Horowitz and Carson (1990), Cropper, Aydede, and Portney (1992, 1994), Johannesson and Johansson (1997), and Frederick (2003).

account for some of the heterogeneity in time preference, and Section 6 summarizes our results and their implications.

## 2. Choice Task and Sample Description

Survey Structure

Our study uses an original survey in which each respondent considered a choice task such as that presented in Figure 1. The general research strategy was to elicit respondents' valuations of improvements that would begin after different periods of delay. Respondents make five choices among three policy options, where each is defined on three dimensions: year when improvement begins, amount of water quality improvement, and cost of the policy per year. Before considering the choices, respondents receive detailed information regarding the water quality and cost dimensions of the choices. In the choice task, the respondent indicates the most preferred choice among the different policy alternatives. The first dimension in Figure 1 is the year when improvement begins, which we will refer to below as time delay. The amount of time delay is zero, two, four, or six years. The second dimension is the amount of water quality improvement, which is the percentage of lakes and rivers in the respondent's region that the U.S. Environmental Protection Agency (EPA) rates as being "good" for fishing, swimming, and quality of the aquatic environment.[6] The percentage improvement ranged from 5 percent to 20 percent. Finally, each of the policies generates costs ranging from \$100 to \$400. The choice design was generated using a structure in which alternatives were balanced with respect to utility (Huber and Zwerina 1996). This

[6] The survey included an extensive discussion of water quality based on the approach taken by the U.S. EPA (1994) in its National Water Quality Inventory. See Viscusi, Huber, and Bell (2004) for further description.

approach increases the amount of information that the choices provide about the parameter estimates.

The structure of the survey design makes it possible to identify empirically the individual's rate of time preference. In effect, for each different time delay, the survey structure permits an estimate of how much the respondent would have been willing to give up in terms of lower water quality or higher cost to remain just as well off. The cost dimension of the policy choice is not needed to estimate this intertemporal tradeoff rate. However, inclusion of the cost component increases the realism of the policy choice and will lead to estimates of the cost-water quality improvement tradeoff that can be compared with estimates using a different survey methodology as an additional validity check on the survey.

The policy choice decisions that respondents faced involved four different levels of cost, four different levels of water quality improvement, and four different periods of time delay, one of which was no delay. We achieve identification of the time delay water quality improvement tradeoff using this formulation by varying the time delay for different levels of water quality improvement.[7]

Modeling the Effect of Delay

The costs and water quality improvements had comparable time dimensions, with each lasting for five years. However, costs uniformly begin immediately while the benefits begin after 0-6 years. Thus the time discounting considers the value in present dollars of the level of improvement or of having the improvement occur sooner. To see the relationship between the present value of costs and improvements, consider the

[7] As an identification check, we also estimated a variety of linear and quadratic specifications and found that the results were robust. See Rust (1994) for further discussion of identification issues.

standard exponential discounting case. There is a delay of t years before the improvement begins. Let the person's utility function be additively separable and linear in cost *c* and water quality improvement *w*, and let the time period of delay be t.[8] Then the present value of the five year imposition of costs beginning immediately is $c[1+\delta+\delta^2+\delta^3+\delta^4]$. Similarly, the present value of water quality benefits after a t year delay is given by $w\delta^t[1+\delta+\delta^2+\delta^3+\delta^4]$. Since the bracketed terms are identical, the person's decision reduces to ascertaining whether the value of c is greater than $w\delta^t$. Put somewhat differently, the cost imposition will be worthwhile if the utility of the water quality improvement in a given year is at least as great as $(1 + r)^t$ multiplied by the utility of the annual cost. Similarly, $1/(1 + r)^t$ units of water quality that will result from improvements begun immediately will be equivalent to a unit of water quality improvement begun after a period of t years. In each instance, the fact that the costs and improvements occur over a five-year period drops out of the analysis, as it serves to establish a policy context and to put costs and time durations of water quality improvements on a comparable basis.

Matters become a bit more complicated based on the quasi-hyperbolic discounting model. The present value of the cost stream becomes $c[1+\lambda\delta+\lambda\delta^2+\lambda\delta^3+\lambda\delta^4]$. If the benefits begin immediately, the present value is $w[1+\lambda\delta+\lambda\delta^2+\lambda\delta^3+\lambda\delta^4]$. The policy is attractive if the utility of the annual water quality improvement w exceeds the disutility of the annual cost. If there is a time delay of t years, benefits are $w\lambda\delta^t[1+\delta+\delta^2+\delta^3+\delta^4]$. The bracketed expression and the λ term are present for all nonzero periods of delay.

[8] We also assume that the discount rate is the same for costs and for improvements. This assumption facilitates the theoretical discussion and is the norm in the literature, but it is not essential for the interpretation of the empirical results because the cost time stream never varies.

Consider the five year stream of water quality improvement deferred by t years that is equivalent to the disutility of the five year cost stream that begins immediately. Let costs be multiplied by –1 to reflect that fact cost c has a negative utility value. The value of w must satisfy

$$w\delta^{t}\left[\lambda+\lambda\delta+\lambda\delta^{2}+\lambda\delta^{3}+\lambda\delta^{4}\right]=-c\left[1+\lambda\delta+\lambda\delta^{2}+\lambda\delta^{3}+\lambda\delta^{4}\right] \tag{2}$$

or

$$w=\frac{-c\left[1+\lambda\delta+\lambda\delta^{2}+\lambda\delta^{3}+\lambda\delta^{4}\right]}{\delta^{t}\left[\lambda+\lambda\delta+\lambda\delta^{2}+\lambda\delta^{3}+\lambda\delta^{4}\right]}>\frac{-c}{\delta^{t}}. \tag{3}$$

Compared to the exponential discounting case, hyperbolic discounting boosts the water quality improvement needed to achieve indifference with the utility of the immediate cost stream when compared to the exponential discounting case. This relationship reflects the general phenomenon that hyperbolic discounting differentially reduces the value of all deferred payoffs by a multiplicative parameter $\lambda$ in the quasi-hyperbolic discounting model. The discussion below will use the exponential discounting case as the reference point and then examine whether the findings are more consistent with that approach or hyperbolic discounting.

<u>Sample Characteristics</u>

In 2004 a group of almost three thousand respondents participated in our valuation survey. The sample participants were members of the Knowledge Networks panel, which is a nationally representative computer-based sample. People who do not have computers are given free internet access through a WEB TV device so that the panel composition closely parallels the U.S. Census statistics. Sample participants were compensated for their participation. The response rate to our survey from members of the panel who were offered an opportunity to participate in the study was over 75 percent. As documented in Appendix Table A, the demographic profile of our respondent group is very similar to the mix of the age 18 and over U.S. population.[9] We describe the properties of the sample and present tests of the survey methodology elsewhere.[10]

Although the survey is not a contingent valuation survey, it is in the general family of stated preference surveys, so that it is essential to provide validity tests for the responses.[11] Chief among these tests to be examined below will be a series of scope tests to ascertain whether subjects prefer more water quality improvement to less and, similarly, lower values of costs and shorter delays are preferred to higher costs and longer delays.[12] In addition, the survey included an additional series of rationality tests to determine whether subjects made decisions that did not lead to the choice of a dominated alternative. Overall, 95 percent of the original sample, or 2,914 individual respondents, passed the dominated choice test and will constitute the sample considered here.[13]

The computer-based survey lasted an average of 25 minutes and included detailed information pertaining to the meaning of water quality ratings and financial costs so as to engage people in the survey task. Each of the 2,914 respondents considered a series of five policy choice tasks, such as that in Figure 1, so that there are a total of 14,570 decisions among the three policies.

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[9] The findings in Viscusi, Huber, and Bell (2004) report sample characteristics and tests for influences such as sample attrition bias and selection effects. That earlier working paper has been augmented by an additional wave of survey respondents for this paper, and this paper also focuses on a different set of survey questions.
[10] Magat, Viscusi, and Huber (2000) and Viscusi, Huber, and Bell (2004) describe these other aspects of preliminary versions of the survey. The current paper provides the first analysis of the questions pertaining to rates of time preference.
[11] Arrow et al. (1993) discuss the importance of rationality tests as a validation check for stated preference surveys.
[12] As emphasized by Heberlein, Wilson, Bishop, and Schaeffer (2005), passing a scope test is instructive, but does not ensure validity.
[13] The empirical estimates reported here are very similar to those obtained using the full sample.

## 3. Empirical Model

Model Structure

To analyze the choice task we employ the random utility model developed McFadden (1974).[14] We begin with a conditional logit model and then explore the sensitivity of the results using a mixed logit formulation. Let $U_{ni}$ be the utility that individual n derives from choice i. The value of $U_{ni}$ consists of the observable portion $V_{ni}$ and the unobservable random component $\varepsilon_{ni}$, or

$$U_{ni} = V_{ni} + \varepsilon_{ni} . \tag{4}$$

The value of the observable component $V_{ni}$ is a function of the policy characteristics $X_{ni}$ and the personal characteristics $Y_{ni}$. For a given choice set, the value of $Y_{ni}$ is the same across all policy choices, i = 1, 2, 3, so we will designate it $Y_n$. We also permit an interaction of the policy attributes $X_{ni}$ and the individual attributes $Y_n$, leading to

$$U_{ni} = X_{ni} + Y_n + X_{ni} Y_n + \varepsilon_{ni} . \tag{5}$$

The probability that individual n chooses policy i is

$$P_{ni} = \Pr ob\left(U_{ni} > U_{nj} \text{ for all } j \neq i\right). \tag{6}$$

Since the $Y_n$ term is common to both $U_{ni}$ and $U_{nj}$ based on the construction of the policy choice task, it plays no role in the estimation, so that the task is to estimate

$$P_{ni} = \Pr ob\left(X_{ni} + X_{ni} Y_n + \varepsilon_{ni} > X_{nj} + X_{nj} Y_n + \varepsilon_{nj} \text{ for all } j \neq i\right). \tag{7}$$

The individual characteristics will enter through the interactions with the three dimensions of policy choice, or

$$V_{ni} = X_{ni} + X_{ni} Y_n . \tag{8}$$

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[14] The notation and formulation that we use here more closely follows that of Train (2003), with modifications for our survey structure.

The conditional logit probability that individual n chooses policy i is

$$\mathrm{Prob}_{ni} = \frac{e^{V_{ni}}}{\sum_{j=1}^{3} e^{V_{nj}}}. \tag{9}$$

We modify this conditional logit formulation to take into account the fixed effects of each particular policy choice set k given to respondents.[15] Let $\psi_k$ denote the choice task-specific intercept. If we modify the model to include this parameter, the general structure of the model is

$$\mathrm{Prob}_{nik} = \frac{e^{\psi_k + V_{nik}}}{\sum_{j=1}^{3} e^{\psi_k + V_{njk}}}, \tag{10}$$

but no explicit parameter estimates for $\psi_k$ are generated.

The conditional logit model imposes two key assumptions. The most restrictive assumption is that the random components within each subject are not correlated, which leads to the assumption of independence of irrelevant alternatives. A second key assumption is that the variation in preferences is captured by observed respondent characteristics included in the model. Below we present comparative results based on a mixed logit framework that relaxes the independence of irrelevant alternatives assumption and permits there to be unobserved heterogeneity in tastes.

Scope Tests

The first set of empirical estimates to be explored is the basic model that includes only main effects. These estimates will be informative in indicating whether higher cost levels and longer delays are negatively valued and larger improvements are positively valued. Thus, the formulation is

[15] Chamberlain (1980) developed the fixed effects framework for qualitative choice models.

$$U_{ni} = X_{ni} = \alpha w_{ni} + \beta t_{ni} + \gamma c_{ni}, \tag{11}$$

where $w_{ni}$ is the water quality improvement, $t_{ni}$ is the time delay before improvement, $c_{ni}$ is the cost, and all these values are for person n and alternative i. Table 1 presents two sets of regression estimates for two different samples, where the first sample considers the responses only to the initial conjoint question and the full sample includes five observations per respondent. The conditional logit estimates for Question 1 include only a single observation for each respondent and thus constitute a more rigorous across-subjects scope test. The estimates for the full sample include observations both across and within subjects and are conditional fixed effects estimates. The first set of regression results in each instance is for the continuous versions of the policy choice variables. The coefficients have the expected signs with more water quality improvements raising the probability that the alternative is chosen, whereas there is a negative effect of both delay and cost. The magnitudes of the effects are also very similar for both Question 1 and the full sample. In each case, all coefficients are statistically significant at the 99 percent level, two-tailed test. The second set of regression estimates for each of the two samples focuses on the discrete form of each of the policy choice variables using dummy variables for three of the four possible variable values. In addition to exhibiting the hypothesized signs, the magnitudes of the variables follow the expected pattern, as larger water quality improvements are increasingly valued and longer delays and higher cost levels become increasingly unattractive.

These results can also be used to derive the willingness to pay for water quality. Taking the total derivative of utility and setting it equal to zero yields

$$dU = \alpha dw + \beta dt + \gamma dc = 0. \tag{12}$$

The marginal value of each unit increase in water quality is given by the marginal rate of substitution between c and w, or

$$\frac{\partial c}{\partial w} = \frac{-\alpha}{\gamma}, \tag{13}$$

which is \$24.96 for the Question 1 estimates and \$23.17 for the full sample.[16] Because our interest in the Question 1 sample is only from the standpoint of an across-subjects scope test, the subsequent analysis focuses on the full sample.

As a first step toward estimating the rate of discount, consider the overall tradeoff between improvement and delay. This marginal tradeoff rate is given by

$$\frac{\partial w}{\partial t} = \frac{-\beta}{\alpha}, \tag{14}$$

which is 2.235 for the Question 1 estimates and 2.186 for the full sample. For the midpoint survey structure water quality improvement level of 12.5 percent, the equivalent water quality with one year of delay based on the full sample estimates satisfies

$$12.5 = \frac{12.5 + 2.186}{1 + r}, \tag{15}$$

where solving for r yields an average value of r of 17.49 percent. The analogous result for the Question 1 responses is 17.88 percent. These estimates of the discount rate are drawn from an oversimplified model that does not permit possible interactions between time delays and improvements, which from an economic standpoint is the central concern and which will be the focus of the analysis below.

[16] These values are very similar to the estimates generated with a different survey methodology reported in Viscusi, Huber, and Bell (2004).

Conditional Logit Estimates of Delay Interactions

Although respondents may have preferences regarding policy delays generally, the main matter of interest is how delays affect their valuation of water quality improvements and what rates of discount are implied by these preferences. To examine these issues, we formulate three conditional logit specifications. The first set of estimates in Table 1 adds a Delay x Improvement interaction term to the main effects equation, or

$$U_{ni} = X_{ni} = \alpha w_{ni} + \beta t_{ni} + \gamma c_{ni} + \theta w_{ni} t_{ni}\,. \tag{16}$$

The utility gain associated with water quality improvements should be smaller for longer delays t so that θ should be negative. The empirical estimates yield the expected negative effect of the interaction of time delay and water quality improvement. Whereas one unit of immediate water quality improvement has a value of 0.1438, the value of an improvement that occurs after one year is (0.1438 – 0.0086) = 0.1352, dropping to (0.1438 - 6 x 0.0086) = 0.0922 by year 6.

This simple interaction constrains the effect of delay to be a constant value of improvement irrespective of the extent of delay. As a result, the marginal effect of long delays on the implied rate of time preference is greater for long delays than for short delays. The implied average rate of discount is 6.4 percent for a one period delay, 6.7 percent for the midpoint delay value of three years, and 7.7 percent for the upper bound delay period of six years. This rising pattern of rates of time preference is the opposite of the hyperbolic discounting pattern, but the observed pattern is a consequence of the constraints imposed on the estimation.

To provide a more realistic picture of how the length of delay affects the discount rate, the second equation estimated in Table 2 includes a quadratic delay interaction with improvement. This specification permits there to be nonlinearity in the influence of delay on the valuation of improvements, leading to

$$U_{ni} = X_{ni} = \alpha w_{ni} + \beta t_{ni} + \gamma c_{ni} + \theta_1 w_{ni} t_{ni} + \theta_2 w_{ni} t_{ni}^{\;2} . \tag{17}$$

The value of $\theta_1$ is negative, and $\theta_2$ is positive, indicating a diminishing effect of delay on the utility of improvements.

The quadratic specification generates the temporal pattern of discounting that is consistent with the hyperbolic discounting model. A one-year delay has an associated rate of time preference of 10.6 percent. This average rate of time preference declines to 10.4 percent for two years, 10.0 percent at the midpoint delay value of three years, 9.7 percent for four years, 9.2 percent for five years, and 8.5 percent for six years. Though these rates of time preference seem high relative to the cost of capital and discount rates used by the government, these estimates are in a more reasonable range than have been found in many studies of real world choices in product markets and the labor market.[17]

Even the quadratic specification imposes some structure on the discount rate variation with the length of the period, as would other continuous models of discounting. Because the delay variable can take on only three nonzero values of two (Delay 2), four (Delay 4), and six (Delay 6) years of delay, we examine a final discrete variable specification for each of these delay periods that captures the full range of these estimates given by

$$U_{ni} = X_{ni} = \alpha w_{ni} + \beta t_{ni} + \gamma c_{ni} + \xi_1 w_{ni} \text{Delay } 2 + \xi_2 w_{ni} \text{Delay } 4 + \xi_3 w_{ni} \text{Delay } 6 . \tag{18}$$

[17] These studies include, for example, the implied discount rates based on appliance energy efficiency decisions, used car purchases, and decisions involving risky jobs. Frederick, Loewenstein, and O'Donoghue (2002) provide a review, and Hausman (1979) provides an early example of this approach.

The introduction of separate categorical variables for the delay periods leads to a more pronounced effect of the initial period of delay, consistent with the hyperbolic discounting model. The results in the final column of Table 2 indicate that the effect on the utility of improvements of a two-year delay is more than half the effect of a six-year delay. The estimates in the final column of Table 2 can also be used to calculate the discount factor, or similarly the implied rate of interest. Consider for the general case of an n year delay, for n = 2, 4, and 6. Since

$$\beta\delta^{n} = \beta + \xi_{n}, \tag{19}$$

the value of $\delta$ is given by

$$\delta = [1 + \xi_{n} / \beta]^{1/n}. \tag{20}$$

The substantial influence of early delays is reflected in the estimated rate of time preference. For a two-year delay, respondents exhibit a 14.3 percent rate of interest. The utility loss associated with a four-year delay is very similar to that of a two-year delay, with the consequence being that the average rate of time preference is 8.4 percent. For the six-year delay, the average implied rate of interest over that period is 8.7 percent, which is also well below the initial value of 14.3 percent.

The value of $\delta$ given by equation 20 is a nonlinear function of the parameters $\beta$ and $\xi_n$. It is nevertheless feasible to construct 95 percent confidence intervals for $\delta$, which are (0.85, 0.89) for a two-year delay, (0.91, 0.94) for a four-year delay, and (0.90, 0.94) for a six-year delay. These estimates for $\delta$ in turn imply confidence intervals for the rates of time preference r, which in terms of interest rate percent, are (11.8, 17.0) for a two-year delay, (6.6, 10.4) for a four-year delay, and (6.7, 10.7) for a six-year delay. The confidence intervals for the four-year and six-year delay are quite similar. The distinctive confidence interval is for the two-year delay, which is reflective of respondents' very different rate of time preference for the more immediate period of delay.

Mixed Logit Estimates

To explore the robustness of the conditional fixed effect logit estimates, we also estimate the equations in Table 2 using a mixed logit model. The mixed logit model generalizes the conditional logit model on several dimensions. First, it does not require the independence of irrelevant alternative assumption of the conditional logit framework. Second, rather than assuming homogeneous preferences and estimating a single parameter for each variable, the mixed logit model yields estimates of the mean and variance of the individual level parameters, thus providing information on the extent of individual variation in the coefficient estimates. Third, the estimation approach we use takes into account unobserved factors that will affect particular policy choices by the respondent, leading to possibly correlated errors across the repeated choices. The utility of person n of policy j in choice set k for the analog of equation 16 is

$$U_{njk} = V_{njk} = \alpha_n w_{njk} + \beta_n t_{njk} + \gamma_n c_{njk} + \theta_n w_{njk} t_{njk} . \qquad (21)$$

Note that compared to the parameter estimates in equation 15 above, the values of $\alpha$, $\beta$, $\gamma$, and $\theta$ are now permitted to vary across individuals in the sample rather than estimating a single value for each parameter. However, for each of the five choice sets, the values of $\alpha_n$, $\beta_n$, $\gamma_n$, and $\theta_n$ are the same across the choices for the particular individual.

The particular estimation approach used is the hierarchical Bayes estimation procedure, which shares the same behavioral model yields estimates virtually equivalent to mixed logit.[18] The hierarchical Bayesian estimation procedure assumes that each individual's parameters come from a mixture of the aggregate distribution of values with choices that the respondent makes. The mixed logit estimation approach assumes that the parameter vector is normally distributed with mean b and covariance W, and the error term $\varepsilon_{njk}$ is iid extreme value. The coefficient vector is assumed to be independent of the stochastic $\varepsilon$ and w, t, and c, which are non-stochastic. In contrast, the hierarchical Bayes procedure treats b and w as stochastic. Both procedures derive their estimates using simulation methods. The approach takes as its prior estimate of the parameters coefficient values that account for the derived heterogeneity across respondents and the individual's choices. Combining the prior with the likelihood function for the data yields the posterior distribution. Gibbs sampling is then used to take repeated measures of b and W from the posterior distribution. Draws are repeated until the conditional posterior estimates converge. As shown in Huber and Train (2001), the hierarchical Bayes estimates are virtually equivalent to those yielded by classical maximum likelihood mixed logit approaches. The mean and variance of the Bayesian estimator are asymptotically equivalent to the classical maximum likelihood estimates. Moreover, the hierarchical Bayes estimation is less subject to problems of identification.

The estimates in Table 3 present the mean value of the estimated coefficients across the sample as well as the standard deviation of the individual coefficients' values. The various coefficients associated with the delay terms have associated standard deviations that are relatively large, indicating quite substantial heterogeneity in rates of discount across the sample.

[18] For discussion of the properties of hierarchical Bayes estimates, see Huber and Train (2001) and Train (2003).

The mean values of the mixed logit parameters indicate tradeoff rates that closely parallel the conditional fixed effect logit results. Because the utility scale is invariant with respect to a positive linear transformation, it is the coefficient ratios and relative coefficient values that are most instructive. The rates of time preference for the linear Delay x Improvement interaction are almost identical to the conditional logit estimates: 6.3 percent for a one-year delay, 6.7 percent for the midpoint survey delay period of three years, and 7.7 percent for the six-year delay period.

The second column of estimates in Table 3 presents the mixed logit version of the policy choice equation in which there is both a linear and quadratic delay interaction. The linear Delay x Improvement interaction is negative, and the quadratic (Delay x Improvement) interaction is positive, as in the conditional logit estimates. The implied rates of time preference are consistent with a hyperbolic discounting pattern of declining rates of interest, with the values dropping from 10.9 percent with a one-year delay to 10.3 percent for a two-year delay, 9.5 percent for a three-year delay, 8.5 percent for a four-year delay, 7.4 percent for a five-year delay, and 6.2 percent for a six-year delay. The higher rates of interest for short delays are very similar to the conditional logit estimates, and there is a declining rate of interest with the extent of delay, but the amount of this decline is greater for the mixed logit estimates of the mean delay-improvement effects across the sample, as compared to the overall sample parameter estimates of the interaction with the conditional logit model.

The differences between the two sets of estimates are apparent in Figure 2, which illustrates the discount factors for improvements occurring with different periods of delay. For the first three years of delay the discount factors are almost identical for the conditional fixed effects logit and mixed logit estimates. Thereafter the discount factor implied by the mixed logit estimates becomes increasingly greater than that implied by the conditional logit model, which is a reflection of the somewhat lower average rate of time preference implied by the mixed logit results.

The final set of estimates in Table 3 presents the unconstrained estimates in which the different delay categories in the survey are interacted with the level of improvement. The implied average rate of interest is 12.7 percent for a two-year delay, 8.0 percent for a four-year delay, and 7.9 percent for a six-year delay. This pattern accords with the hyperbolic discounting model in that there is a very high initial rate of discount followed by a decline and comparative flattening of the rate of time preference. Unlike the conditional logit results, there is no minor increase in the point estimate of the average rate of time preference with a six-year delay.

The unusual pattern of discounting associated with these results can be illustrated by examining the term structure of the implied rates of interest. Let $r_{fg}$ be the implied annual rate of time preference for the time period extending from period f to period g. Consider the estimates using the time delay interval variables for both the conditional logit and mixed logit models. The first column of Table 4 summarizes the pattern of discount factors $\delta_{0g}$ for the three different periods of delay, and column 2 summarizes the average rates of time preference $r_{0g}$. The substantial weight placed on initial payoffs and the decline in average rates of time preference are inconsistent with the standard exponential discounting model. The steepness of the decline in rates of time preference after the initial period generates an additional anomaly in the discounting pattern. Following the literature on term structure of interest rates, one can calculate the discount rate for the marginal two-year period. For the first two-year period, the average rate of time preference and the marginal two-year rate of time preference are 14.3 percent for the conditional logit estimates. The marginal value of the rate of time preference over the period extending from period 2 to period 4 for the conditional logit model is the value of $r_{24}$ that satisfies

$$(1+0.084)^4 = (1+0.143)^2(1+r_{24})^2, \tag{22}$$

or $r_{24}$ = 2.8 percent. Table 4 also reports the marginal rates of time preference for the six-year time delay for the conditional logit results as well as parallel results for the mixed logit estimates. For each set of estimates, the results display a common general pattern. Because the high average rate of interest over different period of delay exhibits a sharp decline and then remains relatively flat, the marginal rate of time preference drops substantially and then increases. To achieve the great drop in the average rate of time preference for four years of delay, the marginal rate must drop substantially. However, because the average rate of time preference declines for a six-year delay but not greatly, the marginal rate of time preference subsequently rises. The high present value premium associated with hyperbolic discounting in this context consequently introduces a lack of monotonicity in the temporal pattern of marginal discount rates.

The pattern of discount factors associated with a g year delay, which is denoted by $\delta_{0g}$, also is anomalous. Based on the empirical structure of the model, $\delta_{00}$ is set equal to 1.0. As shown in Table 4, under the exponential discounting case, the value of $\delta_{02}$ = 0.77 for the conditional logit model and 0.79 for the mixed logit should be the square of their respective average annual discount factors of 0.88 and 0.89. Similarly, if $\delta_{04}$ = 0.72 (conditional logit) or 0.74 (mixed logit), then the associated constant annual value of $\delta$ is given by 0.92 (conditional logit) and 0.93 (mixed logit). These discount factors are above the annual value for the initial two-year delay. Finally, for $\delta_{06} = 0.61$ (conditional logit) and 0.63 (mixed logit), the implied annual value of $\delta$ assuming exponential discounting is 0.92 (conditional logit) and 0.93 (mixed logit). Thus, the implied annual discount factor assuming exponential discounting begins at a low level, then rises and flattens out.

4. The Quasi-Hyperbolic Discounting Parameter λ

Estimates of the Quasi-Hyperbolic Discounting Parameter λ

The analysis thus far has presented estimates of annual discount factors based on the exponential discounting framework. It is clear from the observed discount rate pattern that the results are inconsistent with this formulation and have the general characteristics associated with the hyperbolic discounting model. If we recast the results in the quasi-hyperbolic discounting framework, it is possible to generate estimates of the quasi-hyperbolic discount rate parameter λ that governs the extent of the departure from exponential discounting.

First, consider the implications of the conditional logit results. The utility of a one unit improvement with a two-year delay is $\lambda\delta^2 w\left[1+\delta+\delta^2+\delta^3+\delta^4\right]$, which is 0.114 based on the coefficient estimates. Similarly, the utility of a unit improvement with a four-year delay is $\lambda\delta^4 w[1+\delta+\delta^2+\delta^3+\delta^4]$, which is 0.107. The ratio of these utilities is $\delta^2$, which produces an estimate of $\delta$ of 0.969. Taking the ratio of the zero delay utility to the utility after a two-year delay produces

$$\frac{0.148}{0.114} = \frac{w\left[1+\lambda\delta+\lambda\delta^2+\lambda\delta^3+\lambda\delta^4\right]}{\lambda\delta^2 w\left[1+\delta+\delta^2+\delta^3+\delta^4\right]}. \qquad (23)$$

After substituting for the value of $\delta$ of 0.969, this equation yields a value of $\lambda$ of 0.48.[19]

Analogously, one could have used the estimate of $\delta$ implied by the ratio of the six-year delayed improvement to the two-year delay. Because of the change in rates of time preference over time, this approach yields a somewhat different estimate of $\delta$ of 0.943, which implies a value of $\lambda$ of 0.58.

One can generate similar estimates based on the mixed logit results. Using the two-year and four-year delay results we generate a value of $\delta$ of 0.967 and a value of $\lambda$ of 0.53. With the six-year versus two-year delay as the initial contrast, $\delta = 0.948$ and $\lambda = 0.61$. The final column of Table 4 summarizes these results.

Both the conditional logit and mixed logit estimates reflect a similar pattern, with $\lambda$ ranging from 0.48 to 0.61, with the higher values derived from the $\delta$ values based on the longer periods of delay.[20] In each case, however, the discrepancy between the values of $\lambda$ and 1.0 serves as a measure of the extent of departure from the exponential discounting model.

<u>Probabilistic Deferred Benefits</u>

The general structure of the survey provided respondents with a series of hypothetical policy choices affecting water quality in their region, including choices in question blocks in addition to those being considered here. As a result, it would not be reasonable for respondents to assume that all of these policies would be implemented.

[19] This calculation assumes that respondents processed the five year period of water quality improvements, which appears twice in the survey text in Figure 1. Post-survey debriefings of respondents revealed no evidence of misunderstanding of the length of the period of improvement and did indicate explicit awareness of the length of the period.
[20] For interesting results regarding market choices and a review of estimates of the quasi-hyperbolic discounting parameter, which are in the 0.5 to 0.8 range, see DellaVigna and Malmendier (2005).

The possibility that policy outcomes may not actually occur should not be problematic, since there will be no costs incurred for policies that are purely hypothetical. However, if respondents believe that they will be charged for policies offering deferred benefits but that these benefits will not in fact occur, they may exhibit a preference for policies with near term effects if they perceive a greater chance of receiving the benefits. Thus, estimates of rates of time preference may be confounded with time-varying probabilities that the benefits will be received.

This probability issue arises more generally in hyperbolic discounting studies. Recall the standard weighting system of the quasi-hyperbolic discounting case, which we define above as $\lambda\delta^{T}$ for periods T >1 and a value of 1 for immediate payoffs, where $0< \lambda <1$. If respondents believe that the probability of receiving any deferred benefits is some constant value λ, then discounted expected utility with probabilistic benefits and exponential discounting will be equivalent to quasi-hyperbolic discounting with no future uncertainty regarding benefits. If the probability of receiving the benefits λ declines with the delay t, the observed relationship will still adhere to the general hyperbolic discounting relation with declining discount rates over time.

To test for whether respondents believe that there is a negative relationship between time delay and the likelihood of receiving benefits, the survey included a question that asked respondents to assess the probability that the benefits would be received.[21] In an initial large national pretest of several hundred respondents, each

[21] More specifically, the survey wording was of the following form:
*Suppose an improvement in water quality was promised in six years, how likely is it that the improvement would actually occur?*

| | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| *No Chance* | | *Not Likely* | | | *Even Money* | | *Likely* | | *Certain* | |
| *0%* | *10%* | *20%* | *30%* | *40%* | *50%* | *60%* | *70%* | *80%* | *90%* | *100%* |

subject assessed the probability of receiving the benefits for each of the four time delay possibilities, ranging from zero to six years. However, there was very little variation in the assessed probabilities for each time delay period. Such invariance could occur in a within-subject format if respondents anchored on their response for the first time delay period presented and gave the same or very similar assessments for the other periods. To avoid such anchoring effects, the results reported here are for single time delay questions administered across subjects.

Even with this across-subject design there was little variation in the assessed probabilities of receiving the policy benefits. For the zero-year delay case, the probability was 0.373. This assessed value rose to 0.422 for a two-year delay and 0.421 for a four-year delay or a six-year delay. While the assessed probability for the zero-year delay is significantly less than for the two-year delay, the magnitude of the difference is not great.[22] The probabilities for delays of two years and four years are virtually identical, with no statistically significant differences among them.

Overall, there is no evidence of increasing skepticism of policy benefits as the time delay is increased. The time-related pattern of discounting for this study is not attributable to time-dependent probabilistic concerns.

### 5. Conditional Fixed Effect Logit Estimates with Interactions

By including a series of interactions of personal characteristic variables with the main effects in the conditional logit model, it is possible to analyze some explicit determinants of individual heterogeneity in the model. Table 5 reports estimates for six main effects and a set of eight interactions for each of these variables.[23] The inclusion of

[22] In particular, the t statistic is 3.66.
[23] In this formulation, we specify all variables in terms of their deviations from the mean.

such a large set of interactions eliminates the statistical significance of the Delay 2 x Improvement interaction, but otherwise the signs of and significance of the main effects are not influenced by the extensive set of interactions.[24]

A variable of central interest with respect to time preference is individual age, as older respondents will have a shorter period of time during which they can experience the benefits of a policy. The age variable has a negative interaction with delay, indicating that policy delays generally are less attractive as one grows older. Moreover, there is a significant negative effect of age on the Delay 2 x Improvement interaction, which is the critical term that drives the extent of hyperbolic discounting and strong present orientation. This positive relationship of age and rates of time preference is also consistent with the evidence on rates of discount of older people found by Read and Read (2004).[25] Moreover, other studies of valuation of environmental benefits have found a negative effect of age, especially for policies with deferred effects such as those relating to global climate change.

Three of the included variables serve as explicit measures of individual differences in preferences toward the environment: whether the respondent is a member of a major environmental group (Environmentalist), whether the respondent makes use of lakes and rivers for recreational purposes (Visited Water), and the density of lakes in the respondent's state of residence (Lake Density). As one might expect, environmentalists were more willing to incur costs that would finance water quality improvements. Those how have visited water bodies have a higher marginal utility of delayed water quality improvements and are less tolerant of policy delays generally. Interestingly, those who

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[24] Several other interactions were explored but were never statistically significant. These included interactions with self-reported stress levels, smoking status, and $(age)^2$.

[25] Read (2001) provides further examination of the age-discounting relationship.

have visited water bodies have significant positive delay x improvement interactions in all three instances. Those who visit water bodies do not reduce their valuation of improvements in water quality to the same extent as the others, but these users of water are more averse to delays generally. Respondents who live in states with a higher density of lakes have a negative overall valuation of delays.

Race and gender are sometimes influential as well. Black respondents have a greater present orientation in that they are more averse to delays generally but not to the six year delay interacted with improvements, and appear to be less cost sensitive. Female respondents are also less cost sensitive than men but otherwise do exhibit not exhibit any statistically significant differences.

The years of education and household income variable should have similar influences to the extent that education is a proxy for lifetime wealth. Higher levels of education also may also reflect a lower revealed rate of time preference. Better educated respondents have greater unwillingness to incur delays of four years. Those with more income are less sensitive to costs, which will raise the cost-improvement tradeoff that they are willing to make.

## 6. Conclusion

People would rather have improvements in environmental quality that occur now rather than later. This expected result is quite consistent with economic rationality. Similarly, some results such as the positive relationship of education and income on valuations of the environment and the effect of age in accentuating the importance of time delay also followed expected patterns. The survey results also passed consistency tests in that more water quality was preferred to less, and steadily higher levels of cost or time delay were each viewed as increasingly undesirable.

Notwithstanding these plausible aspects of behavior, the intertemporal preferences displayed by respondents were inconsistent with the stationarity property in the canonical exponential discounting model. Revealed annual rates of time preference decline with the time delay before benefits occur. The estimated pattern of discount factors for different periods of delay is consistent with observed patterns in the hyperbolic discounting literature. To accommodate the initial steep decline and subsequent flattening of the rate of time preference-time delay relationship, the temporal pattern of marginal rates of time preference declines and then rises, but remains below the rate of time preference for the initial period of delay. The estimates of the quasi-hyperbolic discounting parameter $\lambda$ with values from 0.48 to 0.61 indicate the substantial weight placed on immediate payoffs and the presence of a considerable departure from the exponential discounting model. This study also was the first to test for whether uncertainty about remote payoffs accounted for a $\lambda$ value below 1.0. There was no evidence of such an influence.

The temporal inconsistency of individual discount rates combines with the low weight that respondents place on deferred environmental improvements to provide a sobering message for policy makers. Most environmental policies do not have immediate benefits, but individuals have a strong preference for immediate results. These preferences in turn will generate political pressures that will foster a present orientation for government policies. One partially reassuring result is that as the delay time is extended, the marginal discount rate declines so that very long-term effects are not increasingly disadvantaged.

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Figure 1
Water Quality Survey Policy Choice[a]

Imagine again that you have recently moved to another region of the country, where water quality is **50% Good**.

Imagine that the government is considering several policies that would temporarily increase water quality **in your region**. Once the policy is in effect, the improvement lasts for five years, then water quality returns to its previous level. Regardless of when the improvement begins, the cost of each begins immediately and continues for five years.

Which of the three policies below would you most prefer?

| | Policy 1 | Policy 2 | Policy 3 |
|---|---|---|---|
| Year When Improvement Begins | Now | 2 Years From Now | 4 Years From Now |
| Amount of Water Improvement | 5% | 10% | 15% |
| Cost of Policy Per Year | $100 | $200 | $300 |
| Which Policy Would You Prefer | Policy 1 * | Policy 2 * | Policy 3 * |

[a] The survey included the following policy variations—Amount of Water Quality Improvement: 5%, 10%, 15%, or 20%; Cost of Policy: $100, $200, $300, or $400; Timing of Improvement: Now, 2 Years From Now, 4 Years From Now, 6 Years From Now.

Figure 2
Discount Factors Predicted by Quadratic Delay Specification

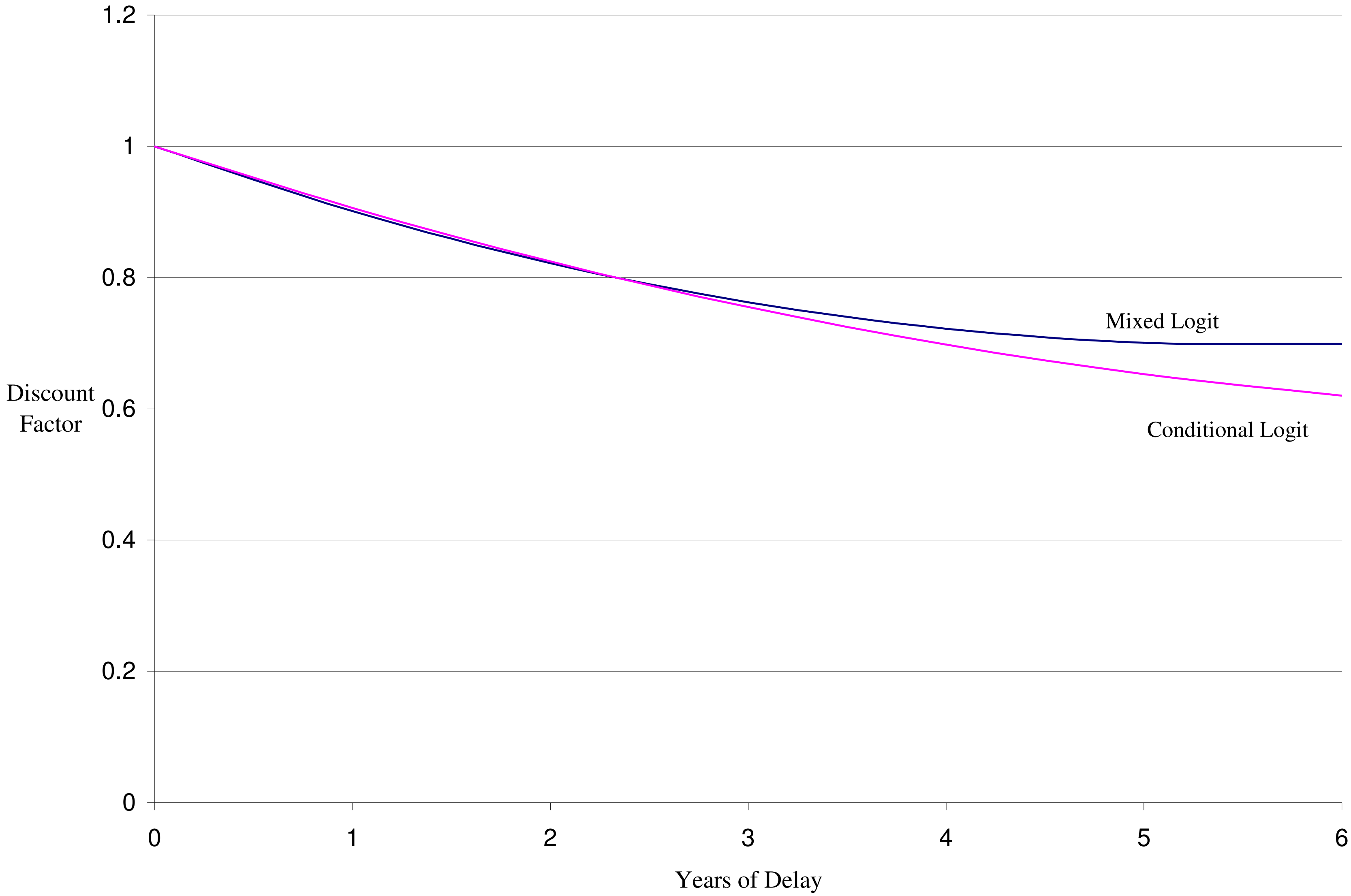

Table 1
Conditional Fixed Effect Logit Estimates of Policy Choice, Scope Test

| | Coefficient (Std. Error) | | | |
|---|---|---|---|---|
| Variable | Question 1 | | Full Sample | |
| Water Quality Improvement 10 | -- | 0.8859***<br>(0.0890) | -- | 0.8545***<br>(0.0352) |
| Water Quality Improvement 15 | -- | 1.2765***<br>(0.1282) | -- | 1.1397***<br>(0.0503) |
| Water Quality Improvement 20 | -- | 1.9108***<br>(0.1168) | -- | 1.8535***<br>(0.0428) |
| Water Quality Improvement | 0.1348***<br>(0.0055) | -- | 0.1205***<br>(0.0022) | -- |
| Delay 2 | -- | -1.0424***<br>(0.0876) | -- | -0.7511***<br>(0.0293) |
| Delay 4 | -- | -1.0745***<br>(0.0981) | -- | -1.0112***<br>(0.0280) |
| Delay 6 | -- | -1.6196***<br>(0.1485) | -- | -1.5486***<br>(0.0403) |
| Delay | -0.3013***<br>(0.0135) | -- | -0.2634***<br>(0.0052) | -- |
| Cost 200 | -- | -0.5374***<br>(0.0996) | -- | -0.5322***<br>(0.0280) |
| Cost 300 | -- | -1.0231***<br>(0.0769) | -- | -1.0814***<br>(0.0285) |
| Cost 400 | -- | -1.4639***<br>(0.1144) | -- | -1.4381***<br>(0.0399) |
| Cost | -0.0054***<br>(0.0003) | -- | -0.0052***<br>(0.0001) | -- |

***Coefficient is statistically significant at the 99 percent level, two-tailed test.

Table 2
Conditional Fixed Effect Logit Estimates of Policy Choice with Delay-Improvement Interactions

| Variable | Coefficient (Std. Error) | | |
|---|---|---|---|
| Water Quality Improvement | 0.1438*** | 0.1472*** | 0.1483*** |
| | (0.0035) | (0.0036) | (0.0036) |
| Delay | -0.1497*** | -0.1339*** | -0.1337*** |
| | (0.0141) | (0.0143) | (0.0144) |
| Cost | -0.0054*** | -0.0053*** | -0.0053*** |
| | (0.0001) | (0.0001) | (0.0001) |
| Delay x Improvement | -0.0086*** | -0.0147*** | -- |
| | (0.0010) | (0.0016) | |
| $(\text{Delay})^2$ x Improvement | -- | 0.0009*** | -- |
| | | (0.0002) | |
| Delay 2 x Improvement | -- | -- | -0.0348*** |
| | | | (0.0031) |
| Delay 4 x Improvement | -- | -- | -0.0410*** |
| | | | (0.0045) |
| Delay 6 x Improvement | -- | -- | -0.0583*** |
| | | | (0.0061) |

***Coefficient is statistically significant at the 99 percent level, two-tailed test.

Table 3
Mixed Logit Estimates of Policy Choice with Delay-Improvement Interactions

| Variable | Mean Coefficient (Std. Deviation of Coefficient) | | |
|---|---|---|---|
| Water Quality Improvement | 0.2236 | 0.2169 | 0.3098 |
| | (0.1385) | (0.1496) | (0.1605) |
| Delay | -0.5015 | -0.5944 | -0.3387 |
| | (0.3781) | (0.5623) | (0.3336) |
| Cost | -0.0098 | -0.0103 | -0.0111 |
| | (0.0067) | (0.0072) | (0.0086) |
| Delay x Improvement | -0.0133 | -0.0235 | -- |
| | (0.0210) | (0.0253) | |
| $(\text{Delay})^2$ x Improvement | -- | 0.0021 | -- |
| | | (0.0065) | |
| Delay 2 x Improvement | -- | -- | -0.0657 |
| | | | (0.0603) |
| Delay 4 x Improvement | -- | -- | -0.0818 |
| | | | (0.0658) |
| Delay 6 x Improvement | -- | | -0.1131 |
| | | | (0.0816) |

Table 4
Average and Marginal Rates of Time Preference

| Time Period | Average Total Discount Factor $\delta_{0g}$ | Average Rates of Time Preference $r_{og}$ | Marginal Rates of Time Preference $r_{g-2,g}$ | Quasi-Hyperbolic Parameter $\lambda$* |
|---|---|---|---|---|
| Conditional Fixed Effect Logit Estimates: | | | | |
| 2 years | 0.77 | 14.3 | 14.3 | -- |
| 4 years | 0.72 | 8.4 | 2.8 | 0.48 |
| 6 years | 0.61 | 8.7 | 9.3 | 0.58 |
| Mixed Logit Estimates: | | | | |
| 2 years | 0.79 | 12.7 | 12.7 | -- |
| 4 years | 0.74 | 8.0 | 3.5 | 0.53 |
| 6 years | 0.63 | 7.9 | 7.7 | 0.61 |

* These $\lambda$ values are calculated using the comparison of the estimates for the years indicated to the estimates for a delay of two years.

Table 5
Conditional Fixed Effect Logit Estimates of Policy Choice with Personal Characteristic Interactions

**Panel A**

| Variable | Coefficient | (Std. Error) |
|---|---|---|
| Water Quality Improvement | 0.0966 *** | (0.0045) |
| Delay | -0.2608 *** | (0.0053) |
| Cost | -0.0055 *** | (0.0001) |
| Delay 2 x Improvement | -0.0079 | (0.0060) |
| Delay 4 x Improvement | -0.0312 *** | (0.0062) |
| Delay 6 x Improvement | -0.0517 *** | (0.0064) |

**Panel B**

| | Improvement Interactions | Delay Interactions | Cost Interactions | Delay 2 x Improvement Interactions | Delay 4 x Improvement Interactions | Delay 6 x Improvement Interactions |
|---|---|---|---|---|---|---|
| Age | -0.0005 * | -0.0024 *** | 8.41e-6 | -0.0007 ** | -0.0004 | 0.0003 |
| | (0.0003) | (0.0003) | (6.53e-6) | (0.0004) | (0.0004) | (0.0004) |
| Black | -0.0099 | -0.0822 *** | 0.0009 *** | 0.0023 | -0.0127 | 0.0571 *** |
| | (0.0134) | (0.0162) | (0.0003) | (0.0177) | (0.0188) | (0.0189) |
| Female | -0.0082 | -0.0157 | 0.0005 ** | -0.0034 | -0.0137 | 0.0034 |
| | (0.0090) | (0.0107) | (0.0002) | (0.0119) | (0.0124) | (0.0127) |
| Education | 40.08e-6 | -0.0013 | -5.68e-5 | -0.0024 | -0.0062 ** | -0.0029 |
| | (0.0018) | (0.0021) | (4.4e-5) | (0.0024) | (0.0025) | (0.0025) |
| Income x 1000 | -3.58e-5 | 0.0001 | 8.60e-6 *** | -0.0003 | -3.5e-5 | -0.0002 |
| | (0.0001) | (0.0001) | (2.99e-6) | (0.0002) | (0.0002) | (0.0002) |
| Environmentalist | -0.0250 | -0.0345 | 0.0012 ** | -0.0303 | -0.0369 | -0.0549 * |
| | (0.0201) | (0.0245) | (0.0005) | (0.0265) | (0.0280) | (0.0289) |
| Visited Water | 0.0570 *** | -0.0383 *** | 0.0001 | 0.0239 * | 0.0475 *** | 0.0459 *** |
| | (0.0097) | (0.0116) | (0.0002) | (0.0130) | (0.0135) | (0.0139) |
| Lake Density | -0.0001 | -0.0010 * | 1.4e-5 | -0.0004 | -0.0005 | -0.0002 |
| | (0.0005) | (0.0006) | (1.2e-5) | (0.0006) | (0.0007) | (0.0007) |

Note: All variables are in terms of deviations with respect to their mean values.
*Coefficient is statistically significant at the 99 percent level (***), 95 percent level (**), or 90 percent level (*), two-tailed test.

Appendix: Table A
Comparison of Sample to the National Adult Population[1]

| Demographic Variable | Survey Participants (n=2914) Percent | US Adult Population Percent |
|---|---|---|
| ***Employment Status (16 years or older)*** | | |
| Employed | 60.4 | 62.3 |
| ***Age*** | | |
| 18 - 24 years old | 13.6 | 13.3 |
| 25 - 34 years old | 20.5 | 18.3 |
| 35 - 44 years old | 19.2 | 20.4 |
| 45 - 54 years old | 18.4 | 18.7 |
| 55 - 64 years old | 11.9 | 12.2 |
| 64 - 74 years old | 11.8 | 8.4 |
| 75 years old or older | 4.6 | 8.1 |
| ***Educational Attainment*** | | |
| Less than HS | 18.8 | 15.4 |
| HS Diploma or higher | 59.3 | 57.4 |
| Bachelor or higher | 21.8 | 27.2 |
| ***Race / Ethnicity*** | | |
| White | 80.0 | 81.9 |
| Black/African-American | 13.3 | 11.8 |
| American Indian or Alaska Native | 1.8 | 0.9 |
| Asian/Pacific Islander/Other | 4.8 | 5.5 |
| ***Race / Ethnicity of Household*** | | |
| Hispanic | 10.2 | 12.1 |
| ***Gender*** | | |
| Male | 50.7 | 48.5 |
| Female | 49.3 | 51.5 |
| ***Marital Status*** | | |
| Married | 56.5 | 58.8 |
| Single (never married) | 26.5 | 24.4 |
| Divorced | 11.7 | 10.2 |
| Widowed | 5.3 | 6.6 |
| ***Household Income (2002)*** | | |
| Less than $15,000 | 15.6 | 16.1 |
| $15,000 to $24,999 | 11.7 | 13.2 |

| | | |
|---|---|---|
| $25,000 to $34,999 | 12.1 | 12.3 |
| $35,000 to $49,999 | 18.8 | 15.1 |
| $50,000 to $74,999 | 17.3 | 18.3 |
| $75,000 or more | 24.5 | 25.1 |

1. *Statistical Abstract of the United States, 2004-5*. 2003 adult population (18 years+), unless otherwise noted.