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Exploring the Beta Model Using Proportional Budget Information in a Contingent Valuation Study

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Abstract

Using a set of random telephone and Internet (web−based) survey samples for a national advisory referendum, we implement Beta models to handle proportional budget information, and allow for consistency in modeling assumptions and the calculation of estimated willingness to pay (WTP). Results indicate significant budget constraint effects and demonstrate the potential for Beta models in handling mental−accounting type information.

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1. Introduction and Background

Despite persistent concerns over the potential for upward hypothetical bias in responses, the survey-based contingent valuation (CV) method remains a commonly used approach for assessing non-market values for changes in public and environment goods (e.g., in benefit-cost analyses and natural resource damage assessments). As one strand of inquiry into the validity of estimated values from CV studies, there has been considerable past research and debate on respondents' ability to pay, the notion of relevant budget or sub-budget constraints, and the effect of budget reminders (e.g., Mitchell and Carson, 1989; Arrow et al., 1993; Kemp and Maxwell, 1993; Willis and Garrod, 1993; Loomis et al., 1994 and 1995; Bockstael, 1995; Whitehead and Blomquist, 1995; Kotchen and Reiling, 1999; Boyle, 2003), and also whether the idea of "mental accounts" or sub-budgets can explain responses under different circumstances (e.g., Magnussen, 1992; Bateman and Langford, 1997). It remains an open question as to what budget constraint information is relevant in different CV settings and circumstances. However, the theory of mental accounting (Thaler 1985 and 1990) suggests that people tend to simplify spending decisions by subdividing a larger budget into various "mental accounts." For example, expenditures are grouped into categories, such as for housing, food, etc., and then spending is sometimes constrained by these implicit or explicit budgets (Thaler, 1999). In a recent related study, Li et al. (2005) found that simply asking two mental-accounting type questions, as a split-sample treatment early in the survey, had a significantly negative effect on respondents' willingness-to-pay (WTP) in a CV study.

However, using mental accounting in a CV study raises practical concerns not fully explored in Li et al (2005), such as how to use the additional information collected in modeling and in estimating willingness-to-pay (WTP) under different constraints. For example, the truncation of the upper bound on the WTP distribution may be relevant; should it be respondents' total income, some articulated subset of income for different activities or something in between like total discretionary income? Hanneman and Kanninen (1996) point out that economic theory requires that an individual's WTP function be bounded by income (although this not always done). Further, Haab and McConnell (1998) suggest that, in both estimating and calculating WTP for an environmental public good, there exist some reasonable, minimum criteria, which are consistent with both consumer preferences and household constraints that one would expect to apply. These would include that WTP should have "an upper bound not greater than income" (Haab and McConnell 1998, p. 218). They argue that there should be no arbitrary truncation in either modeling assumptions for estimation or calculation of WTP. Haab and McConnell propose a new model based on the Beta distribution, which accommodates the use of proportional income information (as might be associated with a mental account or sub-budget). To further explore mental accounting and alternative budget effects, this research note investigates the use the Beta model (Haab and McConnell, 1998) with responses to several levels of proportional income information and a unique mixed-mode CV data set.

2. The Data and Experimental Design

The data used in this research consists of a matching telephone (TEL) and Internet survey (KN) samples that were collected in 2000 (for full details see Berrens et al. 2003 and 2004). Each sample was randomly selected. The TEL sample was selected by random digit dialing. The KN sample was provided by Knowledge Networks, which recruits respondents through random digit dialing and then provides them with Web TV technology and free Internet access in return for participation in weekly surveys. The response rates for TEL and KN samples were 45.6% (with 1,699 respondents completed the survey), and 24.1% (with 2,162 completed) respectively. The survey includes demographic, attitudinal, and knowledge questions as well as the valuation section. The posited program change being valued was the implementation of the Kyoto Protocol to reduce global warming. The payment vehicle was higher energy and gasoline prices, and the elicitation format was a national advisory referendum on U.S. household's support for the ratification of the Kyoto Protocol by the U.S. Senate (Berrens et al. 2003 and 2004). The payment amount (PAY) presented to a respondent was randomly allocated across the same set of values in both samples.

Considerable effort was made to provide objective background information that framed the proposed referendum in a neutral way. For details see Berrens et al. (2004), and a web-version of the survey instrument with randomized treatments at: www.unm.edu/~rberrens/gcc/. The contingent scenario provides summary positions by both supporters and opponents of the Kyoto Protocol, and reminds respondents of their household budget. After being presented with the proposed referendum, and a randomized payment amount, respondents were asked to vote For or Against supporting U.S. Senate ratification of the Kyoto Protocol.

Given a variety of design treatments in the overall study, we only look at two directly comparable samples with and without mental account treatments ($N^{TEL} = 493$, $N^{KN} = 564$). The absence (=0) or presence (=1) of the mental account treatment is identified by an indicator variable, MA. Prior to the valuation section, respondents in the mental accounts treatment (MA=1) received two questions: the first question (QA) asked them about the percentage of their household income left over for optional use (charity, for example); the second question (QB) asked them about the percentage of income available for environmental contributions. Thus, respondents under this treatment are encouraged to construct their mental accounting system into income paid for necessary bills and optional uses (QA), and then, even more restrictively, identify income available for environmental contributions (QB). Average percent of income left over for QA is about 18.6% and 17.3%, while it drops to 1.7% and 2.1 % for QB for the TEL and KN samples, respectively.

Explanatory variables used in the Beta modeling include socioeconomic and Kyoto Protocol — related variables. They include annual household income (INC), ideology (IDEO), and respondents' age (YOUNG), environmental group membership (MEMBER) and low annual household income status (LOWINC), respondent's attitude toward international treaties (INTRTY), respondents' knowledge about greenhouse gases (GRH), the effectiveness of the Kyoto Protocol (CONFID) and the fairness (FAIR) of the Kyoto Protocol (for detailed definitions and the associated sample statistics see Berrens et al. 2003,2004, and Li et al.2005).

3. Modeling Considerations

In the Beta model, the probability of a respondent saying No to a specified payment amount (1), and its corresponding estimate of mean annual household WTP (2) are:

$$prob(No) = prob(\frac{WTP_i}{INC_i} < \frac{PAY_i}{INC_i}) = \int_0^{PAY_i / INC_i} p(z)dz \qquad (1)$$

$$E(WTP) = \overline{INC}\frac{a}{a+b} \qquad (2)$$

where $\overline{INC}$ is the mean of relevant allocable income. Since $0 \le \frac{a}{a+b} \le 1$, the estimated mean *WTP* can be viewed as a proportion of respondents' mean income. *P(z)* is the Beta PDF and $z = WTP_i / INC_i$, and INC_i might be set by the relevant budget share or mental account (such as individual's total income, or discretionary income in QA and QB) for each respondent. Then assume that $\frac{a}{a+b}$ is a function of relevant explanatory variables, i.e., $F(\beta' X_i)$, where $F(\cdot)$ is a random function bound by 0 and 1 (Haab and McConnell, 1998). We use the logistic cumulative distribution function. Solving for *b*, the probability of saying No can be written as:

$$prob(No) = prob(\frac{WTP_i}{INC_i} < \frac{PAY_i}{INC_i}) = BN[\frac{PAY_i}{INC_i}, a, a\frac{1-F(\beta' X_i)}{F(\beta' X_i)}] \qquad (3)$$

where $BN(\cdot)$ is the normalized Beta cumulative distribution function. Then, we can estimate the parameters by maximizing the following log-likelihood function:

$$\log L = \sum [W_i \log(1 - BN(.)) + (1 - W_i) \log BN(.)] \qquad (4)$$

To utilize the mental account treatment and budgeting questions, we vary the INC variable in equations (1) and (2). Specifically, we use the respondent's total monthly household income, discretionary income for optional use (QA), and income available for contribution to environmental causes (QB), in the upper limit of the Beta PDF for the model labeled as UL-I , UL-QA, UL-QB, and in estimating the associated mean WTP ($WTP^{UL\text{-}I}$, $WTP^{UL\text{-}QA}$, $WTP^{UL\text{-}QB}$), respectively. While it would seems intuitive that shrinking the upper limit (UL) will necessarily decrease *E(WTP)*, this is not always true. Since *E(WTP)* has two components (equation (5)), $\overline{INC}$ and $\frac{a}{a+b}$, it is not solely determined by $\overline{INC}$. The change of upper limit will affect the estimate of the shape parameter (*a*) as well as the estimates of covariates.

We employ the Beta modeling strategy in four ways. First, for the respondents who only receive the standard reminder of their monthly income (MA=0), we assume they allocate a proportion ($F(\beta' X_i)$) of their income to support the U.S. Senate ratification of the Kyoto Protocol. So, model UL-I is applied, with $WTP^{UL\text{-}I}$ (MA=0) as some proportion of total income. Second, based on the same assumption, for the respondents who received the mental account treatment (MA=1), we also apply model UL-I, with $WTP^{UL\text{-}I}$ (MA=1) as some proportion of total income. Third, for the sample receiving the mental account treatment, we investigate the effect of the proportion of respondent's monthly household income available after major expenses such as housing, transportation, etc. We assume that respondents use budget constraint QA when they answer the referendum question. Model UL-QA is applied to incorporate this information. $WTP^{UL\text{-}QA}$ is some proportion of respondent's monthly discretionary household income after main expenses. Finally, for the sample receiving the mental account treatment, we evaluate voting and estimated WTP behavior under budget constraint QB. After respondents are encouraged to think about income left over after major expenses, they are also asked about the proportion of this income available for contributions to environmental causes. Using the same assumption, we incorporate this information by employing model UL-QB. $WTP^{UL\text{-}QB}$ is some proportion of this income available for contributions to environmental causes.

4. Hypotheses

Based on likelihood ratio tests for restricted and unrestricted models, previous results confirm the presence of a structural break between mental account treatment groups (0.05 level, Chi-square test) as shown in the Li et al (2005) using applied probit models of WTP. Given the presence of such a result, the focus here is on how researchers might then use this additional budget information in a way that is consistent in assumptions between modeling and calculation of estimated WTP. Specifically, we first focus on whether the use of progressively more restrictive budget information translates into significant differences in mean WTP estimates. That is, focusing on the mental account treatment group (MA=1), we expect the estimated mean $WTP^{UL\text{-}QB}$ to be significantly smaller than the estimated mean $WTP^{UL\text{-}QA}$ and $WTP^{UL\text{-}I}$, respectively; and $WTP^{UL\text{-}QA}$ to be significantly smaller relative to the estimated mean $WTP^{UL\text{-}I}$. We test hypotheses *H1a*, *H1b*, and *H1c*, against the null of no differences.

H1a: $WTP^{UL\text{-}QB} < WTP^{UL\text{-}I}$, *for MA=1.*

H1b: $WTP^{UL\text{-}QB} < WTP^{UL\text{-}QA}$, *for MA=1.*

H1c: $WTP^{UL\text{-}QA} < WTP^{UL\text{-}I}$, *for MA=1.*

We expect that the evidence will support the alternative hypotheses in all cases.

In addition to asking whether the use of the additional budget information has any significant effect on estimated WTP, a further issue is whether the conclusion that might be drawn with respect to the test of treatment would be affected. Thus, the additional budget constraint information can also be used in testing a treatment effect. For example, it is also of interest to test for a mode effect in estimated mean WTP, across the two probability-based survey samples (TEL and KN). There could be several possible sources for a mode difference, and given the growing popularity of Internet surveys (see Berrens et al., 2003) this merits investigation. Expressed formally, the null hypothesis of no effect is tested against three alternative hypotheses:

H2a: Given MA=1, $WTP^{UL\text{-}I}$ *(TEL)* > $WTP^{UL\text{-}I}$ *(KN).*

H2b: Given MA=1, $WTP^{UL\text{-}A}$ *(TEL)* > $WTP^{UL\text{-}A}$ *(KN).*

H2c: Given MA=1, $WTP^{UL\text{-}B}$ *(TEL)* > $WTP^{UL\text{-}B}$ *(KN).*

In all cases, we do not expect to reject the null of no difference (see Li et al. 2004). However, drawing broad inference from the recent Legget et al (2003) results, the contrary conjecture is that self-administered surveys (web-based, Internet in our case) may have lower social desirability bias (and thus lower WTP estimates). Finally, the underlying question is whether there is consistency across *H1* - *H3* in the statistical conclusion a research analyst would make.

5. Results

Estimation results presented in Table 1; there are four Beta models presented (one with MA=0 and UL=I, and three with MA=1 and varying budget constraints) for both the KN and TEL samples. Focusing first on *H1,* all t-test results support *H1a* and *H1b* (0.01 level, one tailed test). The t-test results also support *H1c* in 3 of 4 cases (0.0 1 level, one tailed test). The exception is the KN raw data, where there is no statistical difference. Thus, in general, our results support hypotheses *H1a, H1b, and H1*c. Results indicate that estimated mean annual household WTP decreases significantly as more restrictive budget information is incorporated.

Then, after controlling for the mental account treatment effect (models MA=1), we are interested in testing for a mode effect by comparing the estimated mean WTP between the telephone and the Internet survey samples *(H2a, H2b, H2c)* based on common means for the explanatory variables. After the delta method is used for estimating standard errors on the WTP estimates, pair-wise t-tests on the differences in mean WTP are conducted. In this case, the results are mixed. First, the TEL sample estimate of WTP is statistically larger than KN sample estimate if we compare the estimated WTP using total income in the model ($WTP^{UL\text{-}I}$). In contrast, when comparing pairs of estimated WTP for the models using discretionary income ($WTP^{UL\text{-}QA}$ and $WTP^{UL\text{-}QB}$), the TEL sample estimates are not statistically different from KN sample estimates. That is, results support the null hypothesis of no mode differences, but only if discretionary budget information is utilized. We note that this result under restrictive budget constraints is consistent with the finding of common underlying preferences across Internet and telephone survey samples in Li et al (2004). One caveat is that the latter study implements a formal structural model (Cameron et al. 2000) and uses different samples out of various treatments in the larger database of this research project.

6. Conclusions

To summarize, within the context of a survey-based contingent valuation study, this research note examines the effect of choosing from three different and reasonable alternative budget constraints, and explores the use of flexible Beta models that allow consistency between modeling assumptions and the calculation of WTP. Results indicate that the choice of what constitutes relevant budget information has significant impacts on estimated WTP values. We also find mixed evidence of significant survey mode differences (telephone versus Internet) between estimated mean WTP measures. However, the null of no mode difference cannot be rejected whenever the proportional income information (QA and QB) is used. This result indicates that in addition to estimated WTP effects, conclusions on tests of treatment effects (e.g., a mode test in this case) may be sensitive to how information about relevant budget constraints is used. The additional contribution of this research note is to demonstrate that Beta models can be used to incorporate proportional budget information, and thus allow the testing of various mental accounting and similar two-stage budgeting hypotheses

In closing, it remains an open question as to what budget constraint information is relevant in different CV settings and circumstances, and how this might differ across individual households. We don't pretend to fully answer that question in this study, but we do expect that this will be an important future research issue in understanding responses to CV questions. We have shown how choices of the analyst can have effects on results, and how Beta models can be used to facilitate further lines of inquiry. Finally, we would note that the issue of identifying a relevant sub-budget may extend to modeling decisions for both stated preference and revealed preference approaches (e.g., see Bockstael 1995, p. 667), such as travel cost and random utility modeling.

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Table 1. Estimation Results on Beta Models for TEL and KN Samples

	Telephone Sample				KN Sample			
	MA=0	MA=1			MA=0	MA=1		
Variables	UL-I **N=563**	UL-I **N=493**	UL-QA **N=493**	UL-QB **N=493**	UL-I **N=649**	UL-I **N=564**	UL-QA **N=564**	UL-QB **N=564**
Intercept	-.80** (0.78)	-2.19*** (0.05)	-0.17 (0.24)	0.31* (0.19)	-1.27*** (-0.14)	-2.63*** (-0.09)	-0.42** (-0.19)	0.32** (0.16)
INTRTY	0.46 (0.68)	0.15 (0.92)	0.20 (0.14)	0.28** (0.11)	0.56*** (0.17)	0.66** (0.25)	0.39*** (0.14)	0.27*** (0.11)
CONFID	1.17*** (0.26)	1.44*** (0.17)	0.93*** (0.16)	0.76*** (0.13)	0.84** (0.33)	0.92** (0.44)	0.51*** (0.14)	0.53*** (0.11)
FAIR	0.50 (0.54)	0.74 (0.63)	0.32** (0.13)	0.26** (0.11)	0.43 (0.72)	0.40*** (0.10)	0.17 (0.11)	0.12 (0.09)
IDEO	-0.40*** (0.05)	-0.36*** (0.01)	-0.29*** (0.12)	-0.26** (0.11)	-0.76*** (0.17)	-0.77*** (-0.13)	-0.27** (-0.12)	-0.17* (-0.09)
GRHOUSE	0.41 (0.57)	0.32 (0.59)	0.24* (0.13)	0.16 (0.11)	0.44** (0.21)	0.24 (0.16)	0.26** (0.12)	0.20** (0.09)
YOUNG	-0.07 (0.11)	0.59*** (0.19)	0.27 (0.24)	0.17 (0.21)	-0.15 (0.29)	0.13** (0.06)	0.36* (0.22)	0.25 (0.17)
MEMBER	0.13 (0.47)	0.87** (0.35)	0.49 (0.37)	0.28 (0.32)	0.08 (0.44)	0.51 (0.71)	1.29*** (0.46)	1.00** (0.44)
LOWINC	-0.01 (1.22)	0.82*** (0.07)	0.40 (0.30)	0.29 (0.26)	-0.03 (0.07)	0.10 (0.14)	-0.04 (0.26)	-0.37** (-0.18)
ln(*a*)	-1.44* (0.57)	-1.91*** (0.41)	-3.29*** (0.57)	-3.96*** (0.85)	-2.12*** (-0.67)	-1.73*** (0.47)	-2.94*** (-0.32)	-3.12*** (-0.45)
(-2*LL)	530.3	510.9	510.8	511.6	659.8	606.5	628	655
AIC	550.3	530.9	530.8	531.6	679.8	626.5	648	675
Mean$WTP (SE)	17762 (8066)	8572 (592)	6299 (518)	536 (32.9)	10476 (825)	4181 (519)	5564 (406)	526 (30)

Notes:
1. UL-I, UL-QA, UL-QB refer to the models using the upper limit of integration using the total reported income, the monthly income left after major expenses, and the monthly income left for contributions to environmental causes, respectively.
2. Numbers in parenthesis are the standard errors. *, **, *** indicate significance at 0.10, 0.05 and 0.01 levels, respectively.
3. LL refers to log-likelihood values.
4. AIC refers to Akaike Information Criterion value. It is calculated as *(-2)*LL+2P*: , where P is the number of the estimated parameters.