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Abstract

In this article, we consider starting point bias as a heterogeneous phenomenon, that is, respondents in CV surveys do not anchor in the same way. We study the consequences of a mistaken assumption of homogeneous anchoring for the analysis of the shift effect in multiple-bounded dichotomous choice format, when respondents really have heterogeneous anchoring. We show that the shift effect, generally interpreted as incentive incompatibility or "yea"-saying, can be the spurious outcome of disregarded heterogeneous anchoring.

Keywords: Contingent Valuation, Heterogeneous Anchoring, Starting Point Bias, Incentive Incompatibility, Yea-Saying.

JEL Classification: Q26, C35

I Introduction

Starting point bias in double-bounded, or more generally multiple, dichotomous choice (DC) questions is now a well-documented phenomenon in the contingent valuation (CV) literature. In double-bounded DC format this bias generally means that respondents anchor their second answer on the bid of the first DC question (Herriges and Shogren 1996). It has also been combined with other problems concerning multiple-bounded DC questions, such as framing effects[1] (Flachaire and Hollard 2006), "yea/nea"-saying[2] (Chien et al. 2005) and incentive incompatibility[3] (Whitehead 2002). The effect of these factors, as described in the literature, can be seen as inconsistencies in respondents' answers to the first and second questions. In particular, the last two phenomena are treated as a "shift" in respondents' answers between the first and second response (Alberini et al. 1997).[4] Moreover, this literature generally assumes that starting point bias in CV surveys (whether on its own or combined with other phenomena) is a homogeneous phenomenon that can be captured by a single parameter for the whole sample. This, however, is quite a strong assumption since all respondents are supposed to anchor in the same manner. Because this is not necessarily the case, it appears interesting to explore the consequences of a mistaken assumption of homogeneous anchoring on the other phenomena that are combined with the starting point bias.

In this paper, therefore, we explicitly introduce heterogeneous anchoring in the case of a double-bounded model.[5] We compare this to a situation in which homogeneous anchoring is (wrongly) assumed and any shift effect interpreted in terms of incentive incompatibility or "yea"-saying. Our aim is to explore whether a shift effect really does reflect incentive incompatibility or "yea"-saying, or whether it is not simply the spurious outcome of disregarded heterogeneous anchoring. We do so in two steps. In section II, we provide analytical arguments showing that a mistaken assumption of homogeneous anchoring leads to a non-zero expectancy error term and thus to biased parameter estimates. In section III, we show by means of Monte Carlo simulations that a mistaken assumption of homogeneous anchoring can lead to spurious incentive incompatibility or "yea"-saying.

II Theoretical Model

First consider $W^\star_{i1}$ the unobserved respondent i's true estimate of her/his willingness to pay (WTP), which is defined as follows:

$$W^\star_{i1} = X_i\beta + u_i, \qquad u_i \sim NID(0, \sigma^2) \tag{1}$$

where X_i is a set of explanatory variables which represents respondent i's tastes, u_i are Normally and Independently Distributed (NID) error terms and β and σ are unknown parameters. To uncover respondents' WTP, we consider a double-bounded dichotomous choice questionnaire (Hanemann 1985; Carson 1985).

In a double-bounded mechanism, the WTP of the respondent i is not observed, but we observe her/his answers to a first bid A_{i1} and a second bid A_{i2} (follow-up question). When the respondent answers the first bid (A_{i1}), a censoring rule links the observed answer $W_{i1} = 0$, or 1 (resp. "no" or "yes") to the unobserved variable $W^\star_{i1}$:

$$W_{i1} = 1 \quad \text{if } W^\star_{i1} \geq A_{i1} \quad \text{and } W_{i1} = 0 \quad \text{otherwise} \tag{2}$$

If the respondent answers "yes" to the initial WTP question, s/he is then asked for a follow-up bid A_{i2} which is higher than the initial bid ($A_{i2} > A_{i1}$) and lower otherwise ($A_{i2} > A_{i1}$). Herriges and Shogren (1996) consider that the respondent's answer to the follow-up question is a convex combination of the true WTP and the opening bid such that:

$$\text{Model 1: } W^\star_{i2} = (1-\gamma)W^\star_{i1} + \gamma A_{i1} \tag{3}$$

The parameter γ accounts for the anchoring effect (starting point bias) and lies in the interval $[0, 1]$. Whitehead (2002) proposes adding to the WTP equation a "shift" parameter in order to account for potential incentive incompatibility or "yea"-saying. The respondent's answer to the follow-up question is thus determined by the WTP function:

$$\text{Model 2: } W^\star_{i2} = (1-\gamma)W^\star_{i1} + \gamma A_{i1} + \delta \tag{4}$$

Incentive incompatibility or "yea"-saying exists if the "shift" parameter δ is negative or positive respectively; obviously no effect exists if δ equals zero. It is worth noting that in both Model 1 (Eq. 3) and Model 2 (Eq. 4) the anchoring parameter is constant across the sample, *i.e.* each single respondent anchors in the same way. We consider in this article an alternative model in which respondents differ in the way they anchor their answer on the initial bid. The respondent's answer to the follow-up question is then determined by the WTP function:

$$\text{Model 3: } W^\star_{i2} = (1-\gamma_i)W^\star_{i1} + \gamma_i A_{i1} \tag{5}$$

where γ_i is an individual specific anchoring parameter (starting point bias) which depends on respondent's observed characteristics, *i.e.* $\gamma_i = H(Z_i)$. $H(.)$ is a deterministic function and Z_i a set of explanatory variables which can include a subset of X_i.

An interesting issue is, therefore, the consequences of assuming homogeneous anchoring while the true anchoring process is in fact heterogeneous. The individual anchoring effect can be written as:

$$\gamma_i = \bar{\gamma} + (\bar{\gamma}_i - \bar{\gamma}) + h(Z_i) \tag{6}$$

with $\bar{\gamma} = N^{-1}\sum_i \bar{\gamma}_i$ the sample mean anchoring effect, $\bar{\gamma}_i$ the individual mean anchoring effect and $h(Z_i)$ an error term ($E[h(Z_i)] = 0$).[6]

Using the above decomposition, we can write Model 3 (Eq. 5) as a homogeneous anchoring model:

$$W^\star_{i2} = (1-\bar{\gamma})X_i\beta + \bar{\gamma}A_{i1} + \eta_i \tag{7}$$

with error term

$$\eta_i = u_i[1 - \bar{\gamma}_i - h(Z_i)] - (X_i\beta - A_{i1})(\gamma_i - \bar{\gamma}) \tag{8}$$

when the true anchoring process is heterogeneous. By definition, we have

$$E\Big[u_i[1-\bar{\gamma}_i-h(Z_i)]\Big]=0. \tag{9}$$

However, when covariates that explain anchoring also explain WTP, i.e. $Z_i \cap X_i \neq \emptyset$ and/or Z and X are not orthogonal, we have:

$$E\Big[(X_i\beta-A_{i1})(\gamma_i-\bar{\gamma})\Big]\neq 0 \tag{10}$$

Therefore, the error term η_i has a non-zero expectancy. The consequences are two-fold. First, specifying a homogeneous anchoring model (Model 1) when anchoring is in fact heterogeneous would lead to biased parameter estimates and consequently biased mean WTP. Second, adding a shift parameter to a model assuming homogeneous anchoring (Model 2) would "capture" the bias induced by unaccounted heterogeneous anchoring through its shift parameter. One would then mistakenly conclude that there is incentive incompatibility or "yea"-saying, when in fact only heterogeneous anchoring is suggested by the data.

III Monte Carlo Simulations

In this section we conduct two Monte Carlo experiments. For each experiment, we generate respondents' answers by setting at zero the "shift" parameter and imposing heterogeneous anchoring. One important result shows that if, on the contrary, homogeneous anchoring were assumed, either a significant negative or a positive "shift" parameter would be found.

First consider that respondent i's WTP is generated such that

$$W_{i1}^{\star}=\alpha X_{i1}+\beta_1 X_{i2}+\beta_2 S_i+u_i \tag{11}$$

where X_{i1} is a constant term, X_{i2} is a continuous variable generated using a normal distribution with mean 10 and variance 2, S_i is a dichotomous variable generated from a Bernoulli distribution with values 0 (1) with probabilities 0.75 (0.25) and u are the error terms normally distributed with zero mean and variance $\sigma=150$.

The effect of S_i on the WTP is chosen positive in the first experiment and negative in the second experiment. While β_1 always equals 200, the parameters α and β_2 are defined such that the mean WTP is preserved:

Experiment 1: $\alpha=200$ and $\beta_2=200$
Experiment 2: $\alpha=300$ and $\beta_2=-200$

We study here a simple case where anchoring differs according to the dichotomous variable S_i such that:

$$\gamma_i=\gamma_1 S_i+\gamma_2(1-S_i) \tag{12}$$

with $\gamma_1 = 0.2$ and $\gamma_2 = 0.8$. Such an anchoring function indicates that there are two types of respondents in the sample: some are weakly influenced by the first bid ($S_i = 1$) while others strongly anchor their answers ($S_i = 0$). As S_i is generated by a Bernoulli distribution with probability 0.25, the mean anchoring effect $E(\gamma_i)$ equals 0.65. The initial bid design is such that $A_{i1} = 400$ for the observations $i = 1, \ldots, 200$, $A_{i1} = 450$ for the observations $i = 201, \ldots, 400$ and $A_{i1} = 500$ for the observations $i = 401, \ldots, 600$. The average first bid is 450, that is, the mean WTP. When the respondent answers "yes" to a first bid of 400 (resp. 450 and 500), s/he is proposed a second bid of 450 (resp. 500 and 550). When s/he answers "no", the second bid is 350 (resp. 400 and 450). This choice corresponds to the 15 and 85 percentile of the revised WTP distribution in order to avoid inefficiency problems[7] and to focus on heterogeneous anchoring and shift effect. Based on equations (11) and (12) and Model 3 (Eq. 5), we draw 500 artificial datasets of size $n = 1000$ for each experiment. We then estimate Models 1, 2 and 3 by maximum likelihood.[8] Note that for these two experiments, there is no incentive incompatibility or "yea"-saying, that is $\delta = 0$.

Table 1 presents the parameter estimates for each of the three models and their corresponding mean WTP estimate. For each model, results are given for the two experiments. The values in brackets are the cut-off values for the 95% bootstrapped confidence interval across the 500 simulations. Asterisks indicate that the parameter estimates are significantly different from their corresponding true values.

Let us first focus on the true WTP function parameter estimates. Parameters α and β_1 are not significantly different from their true values in all three models. However, the average parameter estimates of Model 3 precisely match their true values, while this is not the case in Models 1 and 2. This also holds for the standard error parameter estimate σ. Results differ for the parameter estimate associated with the WTP covariate S_i which also explains anchoring. Parameter estimate β_2 is significantly biased in Models 1 and 2 (the bias is upward in experiment 1 and downward in experiment 2), while β_2 is not significantly different from its true value in Model 3 and again closely matches its true value whatever the experiment.

The constant anchoring parameter γ in Models 1 and 2 is nearly the same, 0.75, and significantly differs from the mean anchoring effect, 0.65. This bias is a consequence of misspecification of the homogeneous anchoring models. Parameter estimates γ_1 and γ_2 associated with heterogeneous anchoring correspond to their true values when estimated in Model 3.[9]

Consider now the shift parameter δ estimated in Model 2. In the first experiment, WTP is positively influenced by S_i: the shift parameter is positive and differs significantly from its true value 0. It could thus be concluded that the data shows spurious "yea"-saying. The same line of argument is applicable to the second experiment when the covariate negatively influences the WTP. Here, the shift parameter is negative and significantly different from 0, indicating incentive incompatibility - even though there is none. Given these results, it is possible that the shift parameter in empirical applications does not capture incentive incompatibility or "yea"-saying, but rather the heterogeneity bias induced by the mistaken assumption of homogeneous anchoring.

If we look at the mean WTP estimates, we can arrive at the same conclusion. In Model 1 (constant anchoring but without shift parameter), the mean WTP significantly differs from its true value: upward biased in the first experiment and downward in the second experiment. This does not apply to Model 2. Confidence intervals now include the true mean WTP. We thus argue that the shift parameter δ captures, at least partly, the bias induced by heterogeneous anchoring and therefore "corrects" the mean WTP estimate. It is, however, worth noting that the mean WTP estimates in both experiments are close to the bound of the confidence interval (lower bound when positively correlated and upper bound when negatively correlated). It may be expected that in more complex models (for instance, several explanatory variables which explain both WTP and anchoring) the shift parameter would not adequately correct the heterogeneity bias. Results of Model 3 accounting for heterogeneous anchoring are straightforward: the mean WTP does not differ from its true value and on average precisely matches its true value for both experiments.

IV Conclusion

The model proposed by Herriges and Shogren (1996) provides a convenient way to deal with starting point bias in double-bounded CV questionnaires. When a shift parameter is added as proposed by Whitehead (2002), econometrics can test and thus easily control for both anchoring and incentive incompatibility or "yea"-saying effects. However, both models consider anchoring as a homogeneous phenomenon that affects all respondents identically. In this article, we show that if this assumption is mistakenly applied to data that really indicate heterogeneous anchoring, interpretations and results may be misleading. In particular, the mean WTP can be biased and spurious incentive incompatibility or "yea"-saying can appear. A corollary of the latter result is that a significant shift parameter estimate could indicate the presence of heterogeneous anchoring rather than incentive incompatibility or "yea"-saying. This could have major implications for the interpretation of results in CV field data.

There are, however, simple ways of dealing with heterogeneous anchoring in double-bounded CV questionnaires, such as those presented in this article (see Model 3). For field data it is, of course, important to identify the relevant explanatory variables with respect to anchoring. Flachaire et al. (2006) for example propose a method which borrows tools from social psychology. This method is specifically designed to construct an explanatory variable that can sort individuals with respect to their anchoring patterns. It may be, however, that not all aspects of heterogeneity can be observed and therefore there is unobserved heterogeneity. This will need to be dealt with by switching from close-ended follow-up to open-ended follow-up questions, due to econometric identification constraints (Aprahamian et al. 2004).

Table 1: Monte Carlo parameter estimates for the two experiments (n = 1000)

	True Value		Model 1		Model 2		Model 3	
Parameter	Exp.1	Exp. 2	Exp. 1	Exp. 2	Exp. 1	Exp. 2	Exp. 1	Exp. 2
α	200	300	218.8	303.56	213.30	312.08	199.81	299.86
			[151.9,268.9]	[235.2,358.7]	[148.4,256.7]	[247.2,359.1]	[135.9,271.0]	[230.0,356.2]
β_1	20	20	18.99	18.81	18.59	18.91	19.98	20.01
			[14.0,25.5]	[13.4,26.2]	[14.26,24.0]	[13.8,26.24]	[14.12,25.77]	[13.5,27.4]
β_2	200	-200	268.8⋆	-269.19⋆	213.30⋆	-258.10⋆	200.53	-201.03
			[207.5,344.1]	[-341.1,-213.0]	[204.7,322.8]	[-329.8,-206.0]	[151.6,256.0]	[-252.3,-154.3]
σ	150	150	154.97	155.52	152.07	154.53	149.71	150.35
			[122.5,186.1]	[125.1,191.8]	[127.4,182.0]	[129.78,191.O]	[117.2,183.0]	[119.9,186.1]
γ	-	-	0.752	0.753	0.746	0.750	-	-
			[0.702,0.798]	[0.699,0.798]	[0.705,0.791]	[0.706,0.796]		
γ_1	0.2	0.2	-	-	-	-	0.185	0.189
							[-0.111,0.402]	[-0.082,0.405]
γ_2	0.8	0.8	-	-	-	-	0.798	0.799
							[0.746,0.835]	[0.752,0.839]
δ	0	0	-	-	5.49⋆	-5.36⋆	-	-
					[2.1,8.8]	[-8.8,-2.1]		
Mean WTP	450	450	475.51⋆	424.27⋆	462.77	436.61	449.93	449.63
			[461.1,491.7]	[408.1,437.6]	[446.3,480.4]	[419.8,452.5]	[435.6,462.7]	[436.8,462.1]

⋆: significantly different from the corresponding true value at the 5% level.

Notes

[1]The behavioral responses to ascending vs descending iterative questions differ. More specifically, most of the biases occur in the ascending sequence (DeShazo 2002).

[2]Respondents are more likely to answer "yes" (resp. "no") to the follow-up DC question when they have previously answered "yes" (resp. "no") to the first DC question (see for instance Alberini et al. 1997).

[3]This concept is used in discrete choice surveys to characterize the fact that a subsequent bid acts as a signal and may influence respondents' answers to the subsequent DC questions (Alberini et al. 1997; Whitehead 2002; Carson et al. 2000). Note however that the early CV literature considered an elicitation method to be incentive compatible when the dominant strategy for respondents is to state true willingness to pay (Davis 1964; Randall et al. 1974).

[4]To be precise, if this shift is positive, it is generally interpreted as being "yea"-saying and if negative it can be interpreted as incentive incompatibility or "nea"-saying. In the following, we adopt the incentive incompatibility interpretation when the shift is negative, as usually done in the literature.

[5]For the sake of simplicity, we only consider double-bounded DC questions. However results can easily be extended to multiple-bounded models.

[6]The individual anchoring parameter can be broken down such that:

$$\begin{aligned}
\gamma_i &= E(\gamma_i) + [\gamma_i - E(\gamma_i)] && (13)\\
&= E[H(Z_i)] + [H(Z_i) - E[H(Z_i)]] && (14)\\
&= E[H(Z_i)] + h(Z_i) && (15)\\
&= \bar{\gamma}_i + h(Z_i) && (16)
\end{aligned}$$

with $\bar{\gamma}_i$, the individual mean anchoring effect such that $\bar{\gamma}_i = \bar{\gamma} + (\bar{\gamma}_i - \bar{\gamma})$ and $\sum_i(\bar{\gamma}_i - \bar{\gamma}) = 0$.

[7]Hanemann and Kanninen (1999) consider optimal bid design and make such recommendations based on the initial WTP distribution (admittedly in the absence of starting point bias).

[8]We do so by estimating a multinomial model instead of creating a pseudo panel model as in Whitehead (2002) (See Flachaire et al. (2006) for details on the multinomial estimation procedure). We therefore do not consider the restrictions implicit in the pseudo panel model as discussed by Aadland and Caplan (2004). Nor did we "add" a second error term to the second WTP function (see Whitehead 2004), which would complicate the estimation procedure without providing any insights into the reasoning we develop in this paper. Obviously similar conclusions can be reached when considering a bivariate probit instead of a multinomial probit.

[9]Note that the confidence intervals of γ_1 have a negative lower bound since estimations have been performed without restrictions, to keep things simple. It is however not difficult to reparametrize so as to avoid these negative values.

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