Modeling Starting Point Bias as Unobserved Heterogeneity in Contingent Valuation Surveys: An Application to Air Pollution

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Up to now, the starting point bias has generally been considered to be a homogeneous phenomenon. In this article, we treat anchoring as an unobserved heterogeneous phenomenon. Our contribution is twofold. First, we show analytically and by way of simulations that assuming homogeneous anchoring can be hazardous and lead to misspecifications. Second, we propose an econometric model that starts with a dichotomous question and then uses an open-ended question. We finally apply our model to a contingent valuation survey on air quality. Our results suggest that how anchoring is modeled in empirical studies deserves more attention.

Key words: anchoring, contingent valuation, starting point bias, unobserved heterogeneity

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Contingent valuation (CV) studies are increasingly used to evaluate nonmarket goods and services, environmental goods or health care programs, and thus affect decision making and policy formulation. These studies raise methodological issues that have been extensively documented in Mitchell and Carson (1989), and critically assessed in Hausman (1993). In particular, attention has been paid to biases, such as systematic nonrandom sources of error between the "true" value and the stated value of the respondent.

Willingness to pay (WTP) elicitation is certainly a central issue in the CV literature on these potential discrepancies between stated and "true" WTP values. The two oldest and most widely used formats to elicit WTP are open-ended questions (Davis 1964) and dichotomous (closed-ended or referendum) questions (Randall, Ives, and Eastman 1974). Each of these has pros and cons (Loomis 1990). The National Oceanic and Atmospheric Administration (NOAA) panel recommends the dichotomous choice (DC) introduced by Bishop and Heberlein (1979): *"Would you be willing to contribute (or be taxed) D dollars to cover the cost of . . . "*, where D is a bid proposed to respondents. The main arguments in favor of the DC format are that it mimics the decision-making tasks individuals face in everyday life (Herriges and Shogren 1996) and that it reduces incentives for strategic responses. Further, a question with a *yes* or *no* answer is easier to answer than a question requiring independent quantitative estimation on goods unfamiliar to the respondent. As a consequence, closed-ended question formats have been fairly widely used in CV surveys. However, the DC method reveals little about individuals' WTP and requires large sample sizes to reach a sufficient level of precision (Hanemann and Kanninen 1999).

Therefore, Hanemann (1985) and Carson (1985) proposed adding a follow-up question to improve efficiency of DC questionnaires, known as the double-bounded model. This amounts to asking the respondent for a second bid. More recently, Milon (1989) and Green et al. (1998) propose adding an open-ended question rather than a DC follow-up question since "*it provides far more information on WTP as well as information on plausibility of the response*" (Green et al. 1998, p. 111).

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However, the key disadvantage of starting with a DC question is the influence that the value proposed may have on the individual's answer to the subsequent question: different initial bids may lead to respondents stating different WTP. As a result, different initial bids can lead to different welfare estimates and, consequently, to different public decisions. This initial bid effect is known as the starting point bias (Tversky and Kahneman 1974) or, more generally, the anchoring bias.

This phenomenon is observed in many psychological studies dealing with beliefs about purely objective quantities: the length of the Amazon or the height of the tallest redwood (Tversky and Kahneman 1974), the number of physicians and surgeons listed in the local yellow pages (Wilson et al. 1996), the monthly number of gallons of gasoline used by a car owner, the yearly number of inches of rainfall at the wettest spot on earth (Green et al. 1998), the yearly average mileage traveled by car (O'Conor, Johannesson, and Johansson 1999), etc. The initial bid therefore plays a part in basic cognitive constructs used for decisions under uncertainty. When CV studies elicit monetary values, the initial bid may be considered by the respondent as an indication of the quality of the contingent good. The respondent may therefore think the bid gives an approximate range of the "correct" value of this good. Indeed, "*confronted with a dollar figure in a situation where he is uncertain about an amenity's value, a respondent may regard the proposed amount as conveying an approximate value of the amenity's true value and anchor his WTP amount on the proposed amount*" (Mitchell and Carson 1989).

To deal with this problem, Herriges and Shogren (1996) consider a double-bounded model in which the respondents combine their true WTP with the first bid amount to form a revised WTP. They show that when anchoring occurs, the estimations provided by the usual double-bounded model are largely biased. Their analysis as well as all subsequent analyses on anchoring concludes that this effect must be corrected for, in order to obtain reliable WTP estimates. However, anchoring has up to now always been considered as a process that homogeneously affects respondents: all individuals anchor in the same way.

This is a particularly strong assumption that may lead to misleading results if individuals react differently to the proposed bids, since individual heterogeneity can lead to major misspecifications. The possible existence of individual heterogeneity is mentioned by various authors such as Herriges and Shogren (1996), Lechner, Rozan, and Laisney (2003), or Green et al. (1998): "*one might expect the strongest anchoring effects when primitive beliefs are weak or absent, and the weakest anchoring effects when primitive beliefs are sharply defined*" (Green et al. 1998, p. 95). Although very intuitive, this affirmation is hard to test since "primitive beliefs" is not a clearly defined notion, making it difficult to identify characteristics that explain anchoring.

In this article, we examine the possibility of heterogeneous anchoring when a first DC question is used with an open-ended follow-up. Our contribution regarding this assumption is twofold. First, we show both analytically and by way of Monte Carlo simulations that using a misspecified homogeneous anchoring model when the true anchoring process is heterogeneous leads to biased parameter and mean WTP estimates. Second, in order to solve this problem, we propose an econometric model in which anchoring is heterogeneous and related to unobserved heterogeneity, which basically involves defining the starting point bias as a random variable distributed in the population.

The article is divided into three sections. A starting point bias DC model with an open-ended follow-up question and heterogeneous anchoring is developed and studied in the section "Starting point bias as an unobserved effect." The section "An application" applies this model to a CV survey dealing with air quality improvements. The section "Conclusion" concludes the article.

Starting Point Bias as an Unobserved Effect

To model the starting point bias, we adopt the general framework of Herriges and Shogren (1996), based on the literature on bidding games (Randall, Ives, and Eastman 1974; Bishop and Welsh 1985 for instance). It assumes that an uncertain respondent may consider the starting bid as providing information on the "correct" WTP value. In this case, the respondent combines his true WTP with the starting bid to form a revised WTP that he ends up stating as his WTP. Stated WTP may then be upwardly biased if the starting point is set above true WTP and downwardly biased if the starting point is set below true WTP. This approach covers extreme cases, from no anchoring (i.e., the respondent is not influenced by the initial bid) to complete anchoring (i.e., the respondent completely ignores his true WTP). It has been widely used in the literature

because prior WTP and starting bid are usually combined in quite a simple way: either linear combination or weighted average.

Econometric Models

Let us call W_{i1}^* the unobserved respondent i's true estimate of his WTP, which is defined as follows:

$$(1) \qquad W_{i1}^* = X_i\beta + u_i$$

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. This article focuses on a particular elicitation mechanism to gather information on this true WTP, i.e., a DC question with an open-ended follow-up.

This kind of elicitation design can be econometrically modeled as follows. When the respondent answers the bid (b_i), a censoring rule links the observed answer $W_{i1} = 0$, or 1 (respectively "no" or "yes") to the unobserved variable W_{i1}^*:

$$(2) \qquad W_{i1} = 1 \quad \text{if } W_{i1}^* \geq b_i \quad \text{and } W_{i1} = 0 \quad \text{otherwise.}$$

An open-ended follow-up question that provides more information on individuals' WTP values is then proposed and W_{i2} is the WTP given by respondent i.

This format, like the more widely used double-bounded DC format, may induce a starting point bias. That is, respondents may view the bid b_i as providing information on the "correct" WTP and combine their true WTP, W_{i1}^*, with b_i to form a revised WTP, W_{i2}.[1] Following Herriges and Shogren (1996), we model this revised WTP as a convex combination of the true WTP and the bid:

$$(3) \qquad W_{i2} = (1-\gamma)W_{i1}^* + \gamma b_i, \quad \gamma \in [0,1]$$

where γ is a constant parameter for the whole sample, meaning that anchoring is assumed to be homogeneous. Combining equations (1) and (3) leads to:

$$(4) \qquad W_{i2} = (1-\gamma)X_i\beta + \gamma b_i + u_i(1-\gamma).$$

[1] Some authors consider that anchoring may also influence W_{i1}, the respondent's answer to the first bid (Lechner, Rozan, and Laisney 2003). Although it would be of importance in the double-bounded dichotomous choice (DB–DC) model, it does not matter in our analysis since W_{i1} plays no role.

W_{i2} contains all information required to estimate the set of parameters. This model (referred to as Model I) is easily estimated by maximum likelihood using the following likelihood:

$$(5) \qquad L(W_2;\beta,\sigma,\gamma) = \prod_{i=1}^{N}\left[\frac{1}{(1-\gamma)\sigma}\phi\left[\frac{W_{i2}-X_i\beta(1-\gamma)-b_i\gamma}{(1-\gamma)\sigma}\right]\right]^{(1-I_i)} \times\left[\Phi\left(-\frac{X_i\beta}{\sigma}\right)\right]^{I_i}$$

where $\Phi(\cdot)$ ($\phi[\cdot]$) is the standard normal cumulative (probability) density function and $I_i = 1$ if $W_{i2} = 0$; and 0 otherwise. Null WTP are accounted for by left-censoring the WTP distribution, the second right-hand term, in keeping with McFadden and Leonard (1993). This formulation covers cases from no anchoring ($\gamma = 0$) to complete anchoring ($\gamma = 1$). Although widely used (Whitehead 2002; Lechner, Rozan, and Laisney 2003; Holmes and Boyle 2003), this formulation imposes a homogeneous anchoring, γ.

A common way to relax the homogeneity assumption is to consider the starting point bias as an unobserved random variable of parametric distribution F which has to be estimated. Equation (4) is then:

$$(6) \qquad W_{i2} = (1-\gamma_i)X_i\beta + \gamma_i b_i + u_i(1-\gamma_i), \quad \gamma_i \in [0,1]$$

where the γ_i, conditional on covariates Z_i, are independently distributed according to a distribution $F(\gamma_i|Z_i;\alpha)$. Z_i is a set of explanatory variables that may include covariates X_i of the WTP equation and α the associated parameter vector. The explanatory variables Z_i stand for individual heterogeneity, since it seems relevant to consider that specific variables such as level of education, income, self-confidence, degree of information on the good, or occupational status, may influence the way respondents anchor the starting bid. The corresponding likelihood function is:

$$(7) \qquad L(W_2;\beta,\sigma,\alpha) = \prod_{i=1}^{N}\left[\int_0^1 \frac{1}{(1-\gamma_i)\sigma}\phi\left[\frac{W_{i2}-X_i\beta(1-\gamma_i)-b_i\gamma_i}{(1-\gamma_i)\sigma}\right] \times F(\gamma_i|Z_i;\alpha)\,d\gamma_i\right]^{(1-I_i)}\left[\Phi\left(-\frac{X_i\beta}{\sigma}\right)\right]^{I_i}.$$

In the simplest model (referred to as Model II), Z_i is reduced to a constant term, whereas this is not the case in the model of heterogeneous anchoring with explanatory variables (referred to as Model III). As Models I–III have a left-censored WTP distribution, the predicted true WTP values rely on the estimated parameters $\hat{\beta}$ and the standard error $\hat{\sigma}$ in left-censored models:

$$(8) \qquad \widehat{W_{i1}^{*}} = \Phi\left(\frac{X_i\hat{\beta}}{\hat{\sigma}}\right) X_i\hat{\beta} + \hat{\sigma}\phi\left(\frac{X_i\hat{\beta}}{\hat{\sigma}}\right).$$

Once the different models are described, we need to ascertain the consequences of using a misspecified homogeneous anchoring model, when the true anchoring mechanism is heterogeneous. These consequences are twofold (see Appendix A for the analytical proof). First, heterogeneous anchoring induces a form of heteroscedasticity, which leads to a biased variance estimate, if estimates are based on the homogeneous model while anchoring is heterogeneous. An important consequence is that the mean WTP is biased in left-censored models, since the variance estimate is used to compute it (see equation (8)). Second, when anchoring is conditional on explanatory variables that also explain WTP, or not orthogonal with the WTP covariates, the parameter estimates associated with the WTP equation covariates are also biased. As a consequence, the mean WTP will be biased. The following illustrates these results by way of Monte Carlo simulations.

Misspecification Biases

In this section, we perform a set of Monte Carlo simulations to illustrate how, in a simple case, heterogeneous anchoring can affect estimation of the homogeneous model. We design a set of experiments based on an anchoring effect generated by a Bernoulli distribution conditional on a covariate of the WTP equation. In order to focus on the consequences of heterogeneity, left-censoring is not introduced in the simulations. We confirm our analytical results by showing that both mean WTP and WTP variance are biased when explanatory variables that explain WTP also explain anchoring. Moreover, we show that the bias increases with the level of dependency between the anchoring effect and the WTP covariates.

The simulations are based on the following specifications:

1. We consider that respondent i's WTP is generated such that

   $$(9) \quad W_{i1}^{*} = 200X_{i1} + 200X_{i2} + 200X_{i3} + u_i$$

   where X_1 is a constant term, X_2 is a continuous variable generated using a normal distribution with mean 0.5 and variance 4, X_3 is a dichotomous variable generated from a Bernoulli distribution with values 0 and 1 with probability 0.5, and u are the error terms normally distributed with zero mean and standard deviation $\sigma_0 = 50$.
2. The survey sample size is 600.
3. The bid design is such that $b_i = 300$ for the observations $i = 1, \ldots, 200$, $b_i = 400$ for the observations $i = 201, \ldots, 400$, and $b_i = 500$ for the observations $i = 401, \ldots, 600$. The average bid equals 400, that is the mean WTP since $E(X_1) = 1$, $E(X_2) = 0.5$, $E(X_3) = 0.5$, and $E(u) = 0$.
4. Respondent i's answer to the open-ended follow-up question is generated such that:

   $$(10) \quad W_{i2} = (1 - \gamma_i)W_{i1}^{*} + \gamma_i b_i$$

   where γ_i is a random variable with values 0.2 and 0.8. Its probability distribution is conditionally defined with respect to X_3:

	$X_3 = 0$	$X_3 = 1$
$Pr(\gamma_i = 0.2)$	p	$1 - p$
$Pr(\gamma_i = 0.8)$	$1 - p$	p

   Different values of p allow us to vary the level of dependency between the anchoring phenomenon and the variable X_3. For instance, for $p = 0.5$, the distribution of the anchoring parameter γ_i is independent of X_3. When $p = 1$ or $p = 0$ the dependency between γ_i and X_3 is perfect. The following values of p are studied: 0.5, 0.6, 0.7, 0.8, 0.9, and 1.[2]

First, W_{i1}^{*}, the true WTP, is generated according to equation (9) by drawing X_{i1}, X_{i2}, X_{i3}, and u_i. Second, unobserved anchoring is drawn according to probability p and conditional on X_{i3}. The revised WTP, W_{i2}, is computed using equation (10). In the first step, we estimate by maximum likelihood the homogeneous anchoring model given by equation (5), ignoring the fact that anchoring is heterogeneous in the population. In the second step, we estimate the Bernoulli heterogeneous anchoring model used to generate the data. Table 1 presents

[2] The symmetry of the γ_i distribution makes it unnecessary to study $p = 0$, 0.1, 0.2, 0.3, and 0.4.

Table 1. Monte Carlo Simulations (Variance in Parentheses)

Case	Model		$E[\widehat{W}_1^*]$	X_1 Parameter	X_2 Parameter	X_3 Parameter	Std. Dev. $\hat{\sigma}$
$p = 0.5$	Homogeneous	Mean	400.9	199.23	201.8	200.9	260.4
		Variance	(388.36)	(521.96)	(422.43)	(654.86)	(723.39)
	Heterogeneous	Mean	400.5	200.2	200.0	200.0	49.80
		Variance	(316.33)	(10.02)	(1.06)	(18.90)	(2.30)
$p = 0.6$	Homogeneous	Mean	413.5	210.4	203.0	202.4	261.7
		Variance	(458.68)	(461.96)	(425.08)	(630.17)	(794.25)
	Heterogeneous	Mean	400.6	200.2	200.0	200.0	49.84
		Variance	(317.04)	(9.80)	(1.05)	(17.76)	(2.19)
$p = 0.7$	Homogeneous	Mean	425.2	221.6	203.4	203.0	261.6
		Variance	(439.24)	(410.25)	(406.31)	(628.51)	(766.16)
	Heterogeneous	Mean	400.6	200.2	200.0	199.9	49.84
		Variance	(317.74)	(9.61)	(1.05)	(17.49)	(2.19)
$p = 0.8$	Homogeneous	Mean	436.9	232.5	204.2	204.0	260.5
		Variance	(444.78)	(405.39)	(412.73)	(681.29)	(779.35)
	Heterogeneous	Mean	400.6	200.2	200.0	200.0	49.84
		Variance	(317.81)	(9.57)	(1.06)	(17.27)	(2.14)
$p = 0.9$	Homogeneous	Mean	449.4	244.4	204.7	204.7	258.7
		Variance	(436.11)	(406.37)	(435.47)	(761.16)	(843.21)
	Heterogeneous	Mean	400.6	200.2	200.0	200.0	49.83
		Variance	(316.98)	(9.35)	(1.05)	(17.13)	(2.12)
$p = 1$	Homogeneous	Mean	462.8	254.1	207.5	209.1	259.4
		Variance	(437.12)	(274.08)	(405.06)	(595.57)	(687.77)
	Heterogeneous	Mean	400.3	200.1	200.0	200.1	49.89
		Variance	(335.27)	(9.08)	(1.12)	(18.74)	(2.06)

the mean and the variance of the prior WTP W_{i1}^* as well as the mean and the variance of parameters in equation (9) for 500 simulations and the two models.

First consider the estimations of the homogeneous anchoring model when the anchoring parameter and X_3 are independent ($p = 0.5$). In this case, only the standard deviation of the error term is clearly and highly biased (260.4 while the true value is 50). The parameter estimates associated with the covariates and the constant term in the WTP equation are not biased and consequently the mean WTP matches its true value. However, the variances of all parameter estimates are far greater than for the heterogeneous model. Therefore, misspecification of the homogeneous model leads to large efficiency losses which make statistical inference incorrect.

Second, results significantly differ when the level of dependency increases ($p > 0.5$): the parameter estimates associated with the WTP equation covariates and the standard deviation $\hat{\sigma}$ are both biased. Moreover, the biases of each parameter increase with p, that is, the level of dependency between the anchoring parameter and the explanatory variable X_3. For instance, the parameter estimate for the constant term ranges between 210.4 when $p = 0.6$ and 254.1 when $p = 1$, although its true value is 200. As a consequence, the mean WTP is also biased and the bias increases with the level of dependency: from 413.5 when $p = 0.6$ to 462.8 when $p = 1$. In addition to this, the precision of parameter estimates remains relatively poor.

Finally, it is worth noting that the estimation of the heterogeneous anchoring model leads to unbiased parameter estimates whatever the value of p. Moreover, the accuracy of parameter estimates is clearly greater than for the homogeneous anchoring model.[3]

An Application

Survey Design and Sample

The data used constitute a subsample of an interdisciplinary CV survey originally designed to explore both theoretical and empirical issues related to risks associated with air pollution exposure (Chanel et al. 2004). Specifically, the research focused on the

[3] Simulations that account for left-censoring provide similar conclusions on homogeneous and heterogeneous models. In addition, when $p = 0.5$, the homogeneous model yields neither unbiased parameter estimates nor unbiased mean WTP estimate (see technical appendix Aprahamian, Chanel, and Luchini 2007).

residents of the French Bouches-du-Rhône area (1.8 million inhabitants), where the second largest French city, Marseilles, is located. A sample of this population was asked about WTP to reduce their air pollution exposure, through a three-part telephone survey. The first part determined respondents' risk attitudes, beliefs and knowledge of air pollution, and health status. The second part was devoted to the scenario and the WTP questions (see Appendix B): briefly, each respondent expressed his/her preferences between moving to a city with air pollution levels similar to Marseilles or moving to a city half as polluted where the cost of living was higher. The third part elicited general socioeconomic information.

Each respondent was randomly assigned to one of three initial bid values (respectively FRF[4] 300, 400, 500), representing additional monthly expenditure associated with the choice of living in the less polluted city. There was one DC follow-up question, where the bid was 50, 100, and 200 if the respondent answered "no" to the initial bid and 800, 1,000, and 1,200 if he answered "yes."[5] The elicitation mechanism ended with an open-ended question allowing the respondent to give any amount. The empirical analysis focuses on the responses received from the open-ended question and the anchoring effect from the initial bid.[6]

The survey was conducted by an opinion survey company during July 2000 and July 2001 through computer-assisted telephone interviews using four stratification variables (age, sex, residence, and profession). The sample, restricted to individuals aged eighteen and over and living in Bouches-du-Rhône, was found to be representative of the corresponding population (see table 2 for the sample means and standard deviations of key socioeconomic variables).

Of the initial 1,006 interviews, twenty two (2.19%) exhibited protest responses[7] and eighty-one (8.05%) answered the double-bounded binary choice question but did not respond to the open-ended question. A total of 267 (26.54%) of the remaining 903 respondents that did not answer key questions or whose WTP was surprisingly high in relation to their household income were excluded.[8] The remainder consists of 636 out of 1,006 respondents (63.22% of the initial sample).

Empirical Results

Respondents expressed positive WTP in 538 out of 636 cases (84.59%) and null WTP in the remaining ninety-eight cases (15.41%). The set of explanatory variables is relatively large and contains the standard socioeconomic variables and variables related to beliefs, knowledge, and experience of air pollution.

The distribution F, describing how individuals react to initial bids, has to fulfill two requirements. The first requirement is technical: the support of the random variable has to lie between 0 and 1, since anchoring is defined as a convex combination. The second requirement is theoretical: F has to be flexible enough to cover many possible configurations of starting point bias. On the basis of these two criteria, beta distribution appears to be suitable: its density can be (a)symmetric bell-curved or U-shaped depending on parameter values and includes the uniform distribution as a special case.

The econometric estimates are given in table 3. The first column in table 3 presents the results of Model I, which allows for a homogeneous anchoring effect, i.e., there is one single parameter γ taking into account the effect of the initial bid on respondents' answer. Heterogeneous anchoring is introduced in Models II and III.

First, it is worth noting that the level of income (here, the logarithm) is highly significant and positively related to WTP, whatever the model. Thus, the fact that WTP increases with income is a reassuring result and provides

[4] 1 FRF is equivalent to 0.152 EUR (1 EUR = 1.19 USD in March 2006).

[5] Although the typical double-bounded design is for bids to double after a "yes" response and be halved after a "no" response, we chose to base the initial bids on responses to an open-ended question in a pilot survey of fifty-two participants, whose median was 400 FRF. Follow-up bids were chosen at the twentieth and the eithtieth percentile points to cover between 50% and 75% of the WTP distribution (in accordance with Hanemann and Kanninen's [1999] recommendations on optimal bid design in double-bounded CV data). This proved to be a wise choice, since the median of the open-ended answers was 400 FRF in the survey, and the bids covered 70% of the WTP distribution.

[6] This methodological choice amounts to considering that the second bid has no effect on respondent's answer to the open-ended question. We discuss this choice in the conclusion.

[7] Protest responses are respondents who do not answer the WTP questions or who express null WTP and give a reason in open comments that can be described as protests (for instance, "I do not agree with the principle of paying," "I would not pay since I will only move to live in the country," "I do not agree to pay to move to a less polluted place when I can die tomorrow when crossing the street," or "I do not want to pay because the factories are the major polluters").

[8] The upper limit of the ratio WTP/household income was set at 0.15 and as a result, fifty (4.97%) interviews were removed.

Table 2. Descriptive Statistics (n= 636)

Variable	Description	Mean	Std. Dev.
W_{i1}	Answer "yes" to the first DC question (=1)	0.6289	0.4835
W_{i2}	Answer to the open-ended follow-up question (FRF)	495.4	501.84
ln(*Inc*)	Logarithm of monthly household income (FRF)	9.4299	0.5599
Sex	Gender (Male=1)	0.4811	0.5000
Owner	Status of the occupant of the place of residence (Owner=1)	0.5518	0.4976
KnowAir	Respondent declares good knowledge of air pollution (=1)	0.3145	0.4647
KnowATMO	Respondent knows the ATMO official scale of air pollution levels (=1)	0.0927	0.2903
KnowPOLSAT	Respondent knows the POLSAT is (=1) (actually a fictional pollution index)	0.0487	0.2154
NotAffected	Respondent has personally never felt the effects of air pollution (=1)	0.2106	0.4081
Disease	Respondent suffers from one of the following: allergy, asthma, chronic bronchitis, or migraine (=1)	0.4606	0.4988
Health	Respondent's self-assessed health status (0 to 10)	6.5676	1.9778
Hab	Respondent changes habits during highly polluted days (=1)	0.2405	0.4277
Edu2	Senior high school education level (=1)	0.1981	0.3989
Edu3	University-level education (=1)	0.3474	0.4765
Mars	Respondent lives in Marseilles (=1)	0.4701	0.4994
ObjPol	Monitored air pollution level of the respondent's place of residence (−3 to +2)	−0.3511	1.4246
Age	Respondent age (years)	41.4311	17.4043
NAdult	Number of adults in the household	2.2185	0.9345
NMinors	Number of minors in the household	0.6713	0.9674
NoCar	Respondent's household possesses no car (=1)	0.1195	0.3246
BadQual	Respondent says that the air quality of his/her place of residence is lower than Marseilles (=1)	0.1556	0.3628
Assoc	Respondent is a member of an association (=1)	0.2783	0.4485
Smoker	Respondent smokes more than ten cigarettes a day (=1)	0.2830	0.4508

evidence of validity of the method (Bishop and Woodward 1995).

Turning to the effect of the other significant variables, we note the following coherent results. Coefficients associated with knowing the official scale of air pollution levels (*KnowATMO*), changing behavior during highly polluted days (*Hab*), considering that the air quality of the respondent's place of residence is lower than Marseilles (*BadQual*), and respondents' household possesses no car (*NoCar*) are generally significant and positive. The first two explanatory variables indicate people who are aware of the consequences of air pollution and certainly better informed. Indeed, this confirms the intuition that people who are so concerned about air pollution that they try to reduce their exposure by changing their habits during high pollution periods, or who know the name of the official air pollution scale, are willing to pay more to live in a less polluted city than others. Similarly, for people considering—rightly or wrongly—that the air quality in their place of residence is lower than Marseilles' (*BadQual*), a city half as polluted as Marseilles is considered a major improvement, which explains the higher associated WTP. Possessing no car is likely to indicate these respondents' (12%) greater awareness of the environment or a way of life compatible with not owning a car.

The parameter estimate associated with the number of adults in the household (*NAdult*) is negative, which may be explained by the fact that, *ceteris paribus*, an increase in the number of persons in the household reduces the portion of the budget available for air quality improvement. The effect of *Age* is quadratic, with a maximum between age twenty and thirty three depending on the model: the older the respondent, the lower the WTP. Finally, two

Table 3. Estimation Results for Models I to III

Variables	Model I Hom. Anch.	Model II Het. Anch.	Model III Het. Anch. with Expl.
		Parameter (*p-value*)	
Constant	−6703.993**	−6520.080**	−6458.430**
	< 0.0001	< 0.0001	0.0001
ln(*Inc*)	876.895**	769.693**	760.456**
	< 0.0001	< 0.0001	< 0.0001
Owner	163.488	−18.749	−16.101
	0.2109	0.7779	0.8751
KnowATMO	436.949**	187.958*	194.022
	0.0427	0.0966	0.1311
KnowPOLSAT	−207.939	−308.423**	−276.706
	0.4704	0.0357	0.5476
NotAffected	−169.963	−67.921	−80.668
	0.2768	0.4006	0.4639
Disease	−219.731*	−107.839*	−71.673
	0.0859	0.0975	0.3517
Health	−44.321	−11.114	−17.082
	0.1679	0.4933	0.2824
Hab	282.729*	169.373**	161.703**
	0.0514	0.0234	0.01477
Edu3	226.292	32.126	42.312
	0.1059	0.6461	0.6884
Mars	−232.690	−86.664	−97.296
	0.1285	0.2428	0.3225
ObjPol	−36.314	−41.439	−35.126
	0.4626	0.1130	0.1355
Age	14.730	22.353**	24.996**
	0.4673	0.0374	0.0206
AgeSquare	−0.363	−0.346**	−0.375**
	0.1047	0.0031	0.0038
NAdult	−195.436**	−102.862**	−103.520*
	0.0245	0.0058	0.0802
NMinors	121.721*	23.390	19.608
	0.0740	0.4893	0.6131
NoCar	251.284	223.445**	215.357*
	0.2842	0.0420	0.0710
BadQual	414.115**	235.301**	215.959
	0.0146	0.0066	0.1442
Assoc	186.897	89.609	95.761
	0.1648	0.1981	0.1437
Smoker	−182.604	−116.518*	−161.017**
	0.1843	0.0931	0.0217
Homogeneous anchoring parameter γ	0.681**	-	-
	< 0.0001	-	-
$a_i = \exp(Z_i\alpha_1)$			
Constant	-	−1.308**	2.027
	-	< 0.0001	0.2010
Sex	-	-	0.106
	-	-	0.510
ln(*Inc*)	-	-	−0.352**
	-	-	0.0330
$b_i = \exp(Z_i\alpha_2)$			
Constant	-	−1.508**	−2.732**
	-	< 0.0001	0.0068
Sex	-	-	0.293**
	-	-	0.0102
ln(*Inc*)	-	-	0.117
	-	-	0.2771
Residual standard error $\hat{\sigma}$ (SD)	1,407.9393 (205.8979)	674.1690 (32.31001)	673.7908 (31.4626)
Log-likelihood	−4,111.594	−3,475.401	−3,464.510

Note: The symbols * and ** denote 10% and 5% significance levels, respectively.

variables related to health have a negative effect on WTP: smoking more than ten cigarettes a day (*Smoker*) and suffering from a disease potentially related to air pollution exposure (*Disease*).

The effect of being a *Smoker* may initially appear to be a simple income effect related to the portion of expenditure related to tobacco. However, since air pollution mainly affects respiratory functions, another explanation could be that smokers are less sensitive to improved air quality and to their own health, since they would not stand to benefit fully. Finally, the effect of the *Disease* variable remains unexplained.

Let us now consider the anchoring phenomenon. Model I assumes a homogeneous anchoring which appears to be highly significant ($p < 0.0001$) at 0.68. It is informative to compare this estimate with estimates obtained in other CV studies. In a double-bounded DC model, Herriges and Shogren (1996) obtain $\gamma = 0.36$ for a target population, whereas Pollicino and Maddison (2001) obtain $\gamma = 0.57$. In a similar model that considers that anchoring may also occur before the answer to the first bid, Lechner, Rozan, and Laisney (2003) find $\gamma = 0.63$.

In models with a structural shift standing for incentive incompatibility, Whitehead and Finney (2003) find $\gamma = 0.54$ whereas Whitehead (2002) and Whitehead (2004) find contrasting results on two data sets: no anchoring in the first case and a much higher anchoring parameter $\gamma = 0.96$ in the second one.[9]

Holmes and Boyle (2003) estimate different models allowing for dynamic bid anchoring within an attribute-based CV survey, and obtain anchoring weights ranging from 0.25 to 0.38. Green et al. (1998) obtain marginal effects from the starting point bid of 0.22 and 0.28 when eliciting two WTP's using a DC with an open-ended follow-up question. Finally, Chien, Huang, and Shaw (2005), using a double-bounded model that distinguishes anchoring and "yeah-saying," find $\gamma = 0.65$ for the former and an additional 32% for the latter. Since each valued good is unique, it is not possible to draw final conclusions from the above, but our result appears to be reasonable.[10]

Model I considers anchoring as a homogeneous effect for the whole population. We relax this assumption in Models II and III. Let us consider firstly the distribution of anchoring among respondents estimated in Model II. This distribution is characterized by a beta distribution with parameters $a = \exp(Z_i\alpha_1)$ and $b = \exp(Z_i\alpha_2)$ where Z_i reduces to a constant term (the exponential is used to ensure that a and b are positive). Since the parameters of the distribution are less than one (0.2705, 0.2211), our estimation exhibits an almost symmetrical U-shape with mean 0.55, median 0.642, and variance 0.09. Figure 1 clearly indicates that our sample population is composed of: 25% with no or insignificant anchoring (lower than 0.1), 25% with almost-perfect anchoring (higher than 0.98), and the remainder distributed between these two extremes. Thus using one parameter to describe anchoring misrepresents the true heterogeneous nature of the effect, as 50% of the population exhibits markedly different and extreme anchoring.

The next step is to test whether this heterogeneous anchoring depends on explanatory variables. This is done by inserting, in addition to constant parameters, several explanatory variables as estimators of the beta coefficient (i.e., Z_i in equation (7)). The beta distribution is now characterized by the parameters $a_i = \exp(Z_i\alpha_1)$ and $b_i = \exp(Z_i\alpha_2)$, where α_1 and α_2 have to be estimated. We explore different sets of explanatory variables associated with different assumptions on the anchoring effect: gender (*Sex*), level of education (*Edu2*, *Edu3*), level of income (*Inc*), knowledge of air pollution (*KnowAir*, *KnowATMO*), and a proxy of self-confidence (*KnowPOLSAT*).[11]

The only variables that appear as significant with likelihood ratio (LR) tests are gender and level of income. Results are presented in the last lines of the final column of table 3. The parameter estimate for the level of income is significant for the first parameter of the beta distribution while the parameter estimate associated with gender is significant for the second one. Hence, the assumption of constant parameters in the beta distribution is rejected. This is confirmed by an LR test with Model II as the null hypothesis (*p-value* < 0.001). The effects of these two variables on the distribution are presented in table 4. Note that both parameters of the beta distribution are less than one, which leaves the beta density U-shaped. Moreover, men generally anchor less than women, and according to the

[9] A referee has suggested that this could be due to inappropriate estimations (see Aadland and Caplan 2004).

[10] Note that studies eliciting purely objective quantities lead to anchoring of a similar range, between 0.253 and 0.591 (O'Conor, Johannesson, and Johansson 1999; Green et al. 1998).

[11] This variable relates to the fact that the respondent answers *yes* to the question: "Do you know the POLSAT pollution index?", which is actually fictional.

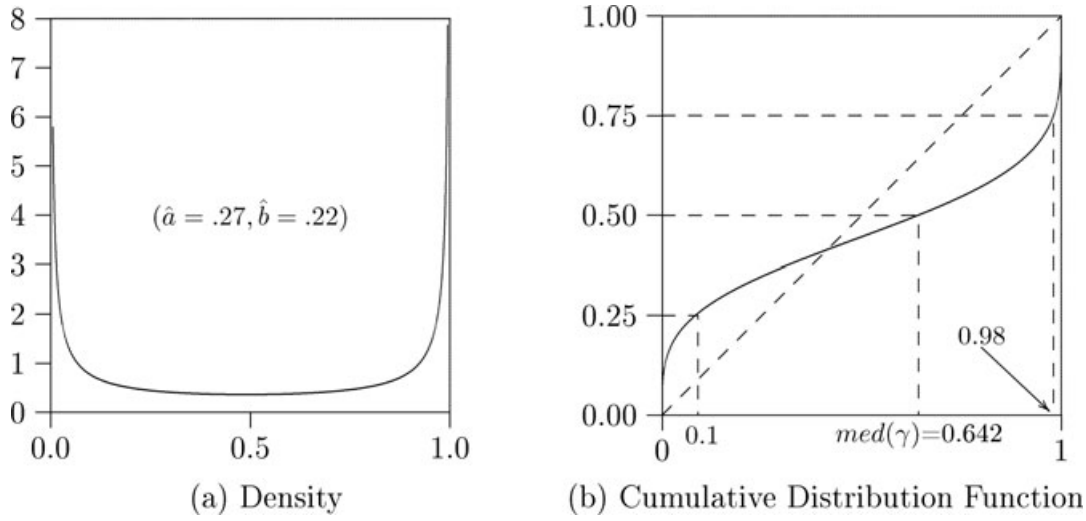

Figure 1. Estimated density and CDF of starting point bias

Table 4. Distribution of Anchoring in Model III: Parameters, Mean, Variance, Skewness, and Quantiles

Gender	Male				Female			
Level of Inc.	Min	Median	Mean	Max	Min	Med	Mean	Max
a	0.5102	0.3050	0.3053	0.2177	0.4588	0.2743	0.2746	0.1958
b	0.2217	0.2630	0.2629	0.2941	0.1653	0.1961	0.1960	0.2193
Mean	0.6970	0.5370	0.5373	0.4254	0.7351	0.5831	0.5834	0.4717
Variance	0.1219	0.1586	0.1585	0.1617	0.1199	0.1653	0.1653	0.1761
Skewness	2.0966	1.9558	1.9558	1.9802	2.2011	1.9911	1.9912	1.9734
$Q_{.25}$	0.4149	0.0922	0.0927	0.0152	0.5017	0.1129	0.1134	0.0164
$Q_{.50}$	0.8781	0.5937	0.5945	0.3027	0.9440	0.7295	0.7302	0.4107
$Q_{.75}$	0.9944	0.9625	0.9627	0.8829	0.9991	0.9905	0.9906	0.9580

quantiles of the distribution, individuals with a high level of income are less likely to anchor than others. Combining these two effects leads to widely differing distribution of the anchoring effect. Figure 2 presents two polar cases by way of illustration: women with the lowest income are much more likely to anchor than men with the highest income (median(γ) = 0.95 versus 0.28).

Estimations and sample characteristics allow us to characterize the distribution of $\widehat{W}_{i1}^{*}$, that is the prediction of the true WTP as defined in equation (8). We used bootstrap procedures to compare the mean WTP and its standard deviation. Briefly, 9,999 replications of size 636 were obtained by drawing with replacement from the 636 $\widehat{W}_{i1}^{*}$ computed for the three models. The resulting empirical distributions of the statistics are convergent and unbiased estimators of the underlying true distributions if both the number of replications and the size of the original data set are large, which is the case according to Efron and Tibshirani (1993). Results for the three models are presented in table 5 and confirm noticeable differences in the parameter estimates and their dispersion.

It is worth noting the discrepancies between mean WTP yielded by the homogeneous model (Model I) and the heterogenous models (Models II and III). The mean WTP for the homogenous anchoring model (Model I) is 1,035 FRF with a bootstrapped standard deviation of 20.3 while the mean WTP for the heterogeneous anchoring models (with and without explanatory variables) are 611.3 and 619.8 FRF with a standard deviation of about 14 FRF. In keeping with our analytical and simulation results, these discrepancies may be explained by two effects. First, the presence of heterogenous

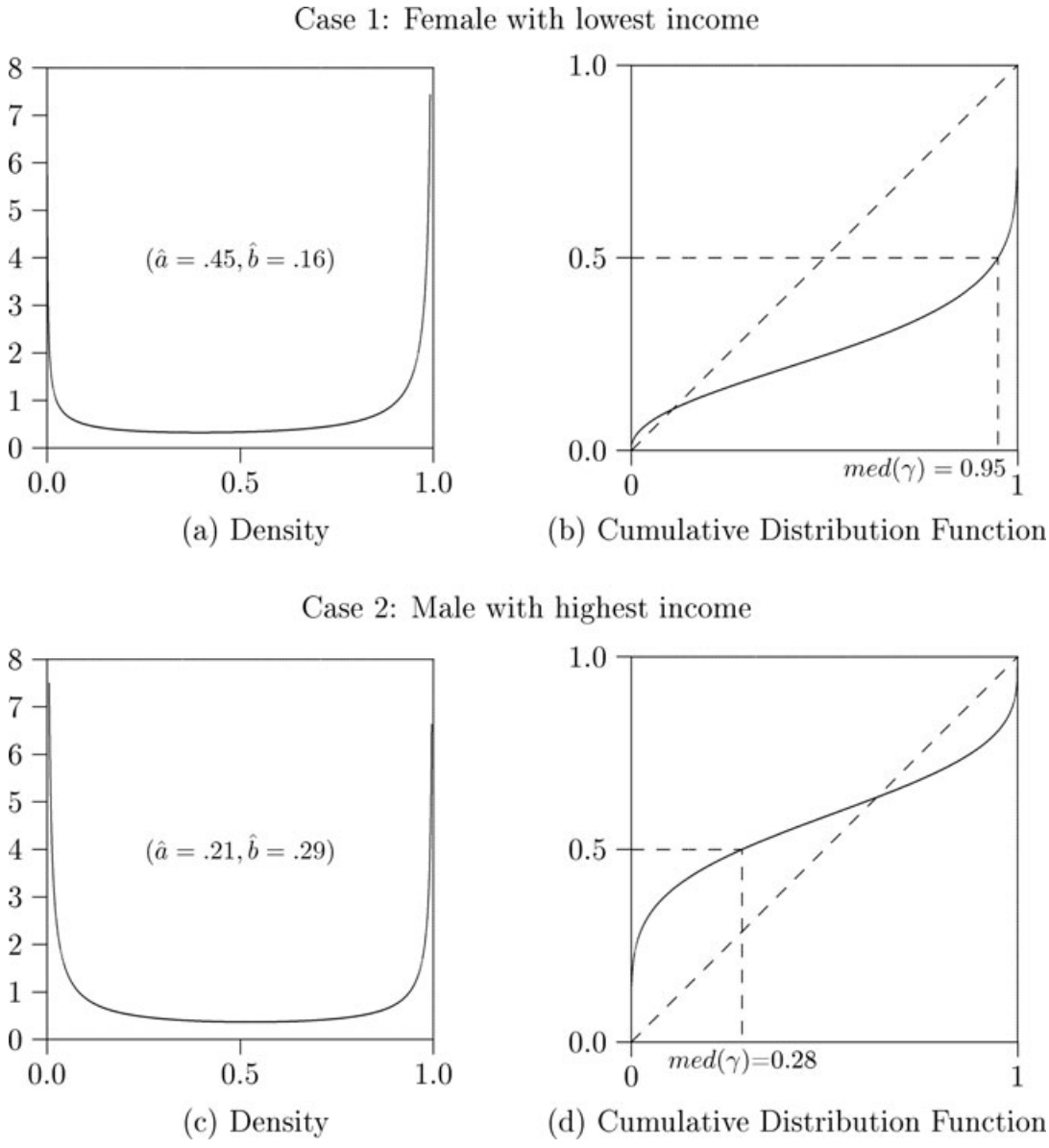

Figure 2. Estimated density and CDF—two illustrative cases

Table 5. Descriptive Statistics for Prior Predicted Bootstrapped Mean $E(\widehat{WTP_{i1}^{*}})$, (9,999 Replications)

Model	Mean	Stand. dev. of the Mean	95% Confidence Interval	Min	Max
Model I:					
Homogeneous anchoring	1,035.25	20.32	[996.22; 1,075.37]	950.75	1,119.94
Model II:					
Heterogeneous anchoring	619.77	13.90	[592.65; 647.12]	561.29	673.08
Model III:					
Heterogeneous anchoring with explanatory variables	611.27	13.78	[584.57; 638.47]	553.76	662.19

anchoring may lead to biased parameter estimates of WTP covariates in the homogeneous model and consequently to a biased mean WTP estimate. Second, a mistaken homogeneous anchoring assumption also induces a form of heteroscedasticity that leads to an upward biased standard error estimate $\hat{\sigma}$. Because in the case of left-censored distribution the mean estimate also relies on the standard error estimate, the biased error term estimate contaminates the mean WTP estimate.

Conclusion

Anchoring has received quite a lot of attention in recent years. Usually, stated preference studies assume that respondents' anchoring is homogeneous. In this article we have analyzed the possibility of heterogenous anchoring in the case of a first DC question with an open-ended follow-up. We used a two-step approach. First, we investigated the nature and extent of the errors that might be caused by mistakenly assuming homogenous anchoring. We have shown that an econometric model based on this mistaken assumption can lead to biased welfare estimates. Second, we have developed an econometric approach in which heterogenous anchoring is considered to be a random variable that stands for both observed and unobserved heterogeneity. Finally, we have illustrated our approach with a CV survey on air quality improvements in France, yielding interesting findings: 25% of respondents either do not anchor or only insignificantly anchor, 25% anchor almost perfectly, and the remainder are distributed between these two extremes.

One limitation of our data must, however, be acknowledged before generalizing the anchoring pattern that we found in our survey, since the elicitation format in the questionnaire introduced a second bid. However, to estimate our econometric model, we assumed that this would not influence the subsequent answer to the open-ended question; without this assumption, econometric identification became problematic. The drawback with this assumption might be that the second bid raises incentive incompatibility problems, i.e., it may act as a signal and influence respondent's answer to the open-ended question (Alberini, Kanninen, and Carson 1997; Whitehead 2002). This incentive incompatibility has been discussed in the context of public good provision. Carson, Groves, and Machina (2000), for example, explain that a second bid functions as if respondents were faced with the same quantity of public good twice but at two different prices and this might induce respondents to answer strategically.

In our application, however, we consider a scenario that provides air quality improvements, usually a public good, privately. Respondents have no influence on the current level of air pollution and can choose between two existing towns with different levels of air pollution and living costs, as in Viscusi, Magat, and Forrest (1988) or Guria et al. (1999). Hence the availability of the good is not influenced by people's actions (or by any institution), as in a public good scenario. Therefore, we can assume that the second DC question will not induce any strategic answers by individuals. While we think that there is no incentive incompatibility problem as such in our application, we note that the second DC might induce certain psychological effects not properly taken into account here. What we do, however, find is heterogenous anchoring.

A consequence of our econometric approach is that it reopens the debate on the relevance of the single DC question with an open-ended follow-up elicitation format initially proposed by Milon (1989) and Green et al. (1998). There are two arguments in favor of this elicitation method: one is that the first DC question is cognitively less difficult to answer and offers the respondent guidance in evaluating the subsequent open-ended question. The other advantage is on a practical, statistical level: an open-ended question gives far more information than a DC question on respondents' evaluations, which typically means that more issues can be identified econometrically. The two main arguments against this elicitation format are that, first, there might be a starting point bias. However, our model provides an econometric tool (heterogenous anchoring as a random variable) that tackles this problem successfully whatever the nature of the anchoring problem for each single individual. Second, there might be strategic incentive incompatibility problems in the case of public good provision. Various authors, however, argue that this problem can be overcome by using appropriate wording (Green et al. 1998) or by informing respondents that the public decision will rely on a quantile (and not on the mean) of the respondents' WTP distribution (Riera 2003). For future research on this question, we propose that one interesting avenue

might be to combine heterogenous anchoring with other phenomena explored in the literature such as "yeah-saying" or the shift effect reflecting incentive incompatibility in the case of double-bounded models (Alberini, Kanninen, and Carson 1997; DeShazo 2002; Whitehead 2002).

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Appendix A: Misspecification Biases: Proof

Consider that the true model is such that anchoring is heterogeneous:

$$W_{i1}^{*} = X_i\beta + u_i \tag{A.1}$$

$$W_{i2} = (1-\gamma_i)W_{i1}^{*} + \gamma_i b_i. \tag{A.2}$$

with u_i an identically independently distributed error term, so

$$W_{i2} = (1-\gamma_i)X_i\beta + \gamma_i b_i + (1-\gamma_i)u_i. \tag{A.3}$$

The individual anchoring effect is generated such that

$$\gamma_i = H(Z_i, \nu_i) \tag{A.4}$$

where Z_i are explanatory variables, ν_i an i.i.d error term and $H(\cdot)$ a deterministic function. Let us assume that the error terms u_i and ν_i are independent:

$$E(u_i.\nu_i) = 0. \tag{A.5}$$

The individual anchoring effect can be written as

$$\gamma_i = E(\gamma_i) + [\gamma_i - E(\gamma_i)] \tag{A.6}$$

$$\begin{aligned} &= E[H(Z_i, \nu_i)] + [H(Z_i, \nu_i) \\ &\quad - E[H(Z_i, \nu_i)]] \end{aligned} \tag{A.7}$$

$$= E[H(Z_i, \nu_i)] + h(Z_i, \nu_i) \tag{A.8}$$

$$= \bar{\gamma}_i + h(Z_i, \nu_i) \tag{A.9}$$

with $E[h(Z_i, \nu_i)] = 0$, which states that the variables Z_i and the error term ν_i are independent, i.e., Z_i are exogenous.

Let us write the individual mean anchoring effect, $\bar{\gamma}_i$ as

$$\bar{\gamma}_i = \bar{\gamma} + (\bar{\gamma}_i - \bar{\gamma}) \tag{A.10}$$

with $\bar{\gamma} = N^{-1}\sum_i \bar{\gamma}_i$ the sample mean anchoring effect and $\sum_i(\bar{\gamma}_i - \bar{\gamma}) = 0$.

The individual anchoring effect, γ_i, is now equal to

$$\gamma_i = \bar{\gamma} + (\bar{\gamma}_i - \bar{\gamma}) + h(Z_i, \nu_i). \tag{A.11}$$

Using this decomposition, the WTP equation is then defined such that

$$\begin{aligned} W_{i2} = &[1 - \bar{\gamma} - (\bar{\gamma}_i - \bar{\gamma}) - h(Z_i, \nu_i)]X_i\beta \\ &+[\bar{\gamma} + (\bar{\gamma}_i - \bar{\gamma}) + h(Z_i, \nu_i)]b_i \\ &+[1 - \bar{\gamma} - (\bar{\gamma}_i - \bar{\gamma}) - h(Z_i, \nu_i)]u_i. \end{aligned} \tag{A.12}$$

This model is equivalent to the following homogeneous anchoring model.

$$W_{i2} = (1-\bar{\gamma})X_i\beta + \bar{\gamma}b_i + \eta_i \tag{A.13}$$

with

(A.14)
$$\eta_i = u_i[1 - \bar{\gamma} - (\bar{\gamma}_i - \bar{\gamma}) - h(Z_i, \nu_i)] - h(Z_i, \nu_i)(X_i\beta - b_i) - (\bar{\gamma}_i - \bar{\gamma})(X_i\beta - b_i).$$

η_i is thus the error term of the homogeneous anchoring model when the true anchoring process is heterogeneous.

First, note that the error term η_i is heteroscedastic and therefore $V(\eta_i) \neq V(u_i)$. Consequently, an estimation of $V(\eta_i)$ is a biased estimation of $V(u_i)$ when one estimates a homogeneous model although anchoring is in fact heterogeneous. This result holds whether the anchoring is a purely unobserved factor or can be explained using observed heterogeneity Z_i.

Second, when the anchoring effect is explained by covariates nonorthogonal to WTP equation covariates, the parameter estimates of the WTP equation are biased. The independence of u_i and ν_i and the exogeneity of Z_i with respect to u_i and ν_i lead to

$$E[u_i[1 - \bar{\gamma} - (\bar{\gamma}_i - \bar{\gamma}) - h(Z_i, \nu_i)]] = 0 \quad \text{(A.15)}$$

and

$$E[h(Z_i, \nu_i)(X_i\beta - b_i)] = 0. \quad \text{(A.16)}$$

However, because the individual mean anchoring effect $\bar{\gamma}_i$ is dependent on the covariates Z_i

$$E[(\bar{\gamma}_i - \bar{\gamma})(X_i\beta - b_i)] \neq 0 \quad \text{(A.17)}$$

when covariates that explain anchoring also explain WTP, i.e., $(Z_i \cap X_i \neq \varnothing)$ and/or Z and X are not orthogonal. In this case, $E(\eta_i) \neq 0$ and the homogeneous model leads to biased parameter estimates and consequently biased mean WTP.

Note that when the anchoring effect is independent of the covariates Z_i, e.g., it is only an unobserved factor:

$$\gamma_i = G(Z_i, \nu_i) = G(\nu_i) \quad \text{(A.18)}$$

the homogeneous model leads to unbiased covariates parameter estimates since $\bar{\gamma}_i$ does not depend upon Z_i, and consequently X_i, through the function $G(\cdot)$.

Appendix B: Hypothetical Scenario and Questions Eliciting WTP

A translation of the questions and the scenario presented to respondents and relevant to this study is reproduced below. Sentences in italics are for the reader and were not read to respondents.

"You are going to be the main actor in our scenario. You will have to take the best decision for yourself and your household.

Let us imagine that you and your household have to move. You can choose between two towns that are exactly equal in terms of inhabitants, work conditions, schools, climate, public services, cultural life, transport, housing, surroundings, etc. There is only one difference between them: the level of atmospheric pollution. The first town—let's call it POL—is as polluted as Marseilles. And the second town—let's call it LESSPOL—is half as polluted as Marseilles.

The problem is that the cost of living is higher in LESSPOL (the less polluted town): housing, local taxes, public transport, etc. are more expensive. This means that if you choose to move to LESSPOL, you will have to pay more to have the same standard of living as in POL.

We would like to know how much you would be willing to pay per month for you and your household to move to LESSPOL (the less polluted town) rather than to POL (the town as polluted as Marseilles). Do not forget that this money will be drawn from your household's budget! You will therefore have less money at the end of the month."

First DC question:
"Would you be willing to pay X French Francs per month for you and your household to move to LESSPOL (the less polluted town) rather than to POL (the town as polluted as Marseilles)?"

Second DC question:
If respondent answers YES: "Would you be willing to pay Y French Francs per month?"
If respondent answers NO or I don't know: "Would you be willing to pay Z French Francs per month?"
with (X, Y, Z) respectively (300, 800, 50), (400, 1,000, 100) and (500, 1,200, 200).

Open-ended question:
"How much would you be willing to pay per month to live in the less polluted town?" ____ French Francs per month.