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Willingness-to-pay in terms of price: an application to organic beef during and after the "mad cow" crisis

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Summary – *This paper aims at assessing consumers' attitudes and willingness to pay for organic beef, an obvious alternative to regular beef in terms of safety, both immediately and at a longer term after the BSE crisis. It is based on two random telephone surveys in an Italian region, the first one conducted in 2001 (few months after the BSE crisis) and the second one in 2003. The analysis is based on an innovative methodology of contingent valuation, keeping into account the possibility for consumers to decide the quantity of their purchase when a price is proposed. The results show that though the effect of the BSE crisis weakened along with time distance, it left some permanent signs in consumers' behaviour. The main conclusion is that the demand for organic beef reduced, but that in the meantime it became more inelastic.*

Keywords: *BSE, organic beef, willingness to pay, contingent valuation*

Consentement à payer en termes de prix : une application à la viande bovine biologique durant et après la crise de la « vache folle »

Résumé – Cet article a pour but d'analyser les attitudes des consommateurs et leur consentement à payer pour la viande bovine biologique (un substitut à la viande bovine conventionnelle) immédiatement à la suite de la crise de la « vache folle » et à plus long terme. Il est basé sur deux enquêtes téléphoniques réalisées par recrutement aléatoire en Italie, l'une en 2001 (peu après la crise de la vache folle), l'autre en 2003. L'analyse est basée sur une méthode innovatrice d'évaluation contingente qui tient compte de la possibilité pour les consommateurs de décider la quantité de leur achat face à un prix proposé. Les résultats montrent que, bien que les effets de la crise de la « vache folle » se soient affaiblis avec le temps, ils ont laissé des signes permanents sur le comportements des consommateurs. La conclusion principale est que la demande de viande bovine biologique a diminuée, mais est devenue moins élastique.

Mots-clés : vache folle, viande bovine biologique, consentement à payer, évaluation contingente

JEL Classification: Q13, Q21

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1. Introduction

The bovine spongiform encephalopathy (BSE) crisis has been one of the most severe food scares in the European food sector. While the first wave mainly affected the UK in 1996, the second wave, in 2001, had deep impacts also in other European countries. Several studies analysed the impact of BSE on demand for beef (*e.g.*, Burton and Young, 1996 and 1997; Burton *et al.*, 1999; Mangen and Burrell, 2001; Pico Ciamarra, 2003; Smith *et al.*, 1999), on the price of beef (Lloyd *et al.*, 2001), on equity prices of meat-related industry (Henson and Mazzocchi, 2002), and on the regional differences in the impact (Caskie *et al.*, 1999). Several of these studies (*e.g.*, Burton *et al.*, 1999; Mazzocchi *et al.*, 2006) suggest that the overall reaction of consumers to the wave of media attention to BSE was a strong immediate reduction in beef consumption and a shift to other meats, followed by a recover. The extent of the recovery is nevertheless disputed, though substantial. For instance, following the discovery of BSE cases in Italy beef consumption underwent a sudden and huge shock with a 49.2% reduction in consumption and a strong shift to other meat, but by the end of 2003 expenditure shares of beef was not far from pre-scare levels, though at lower meat expenditure (Mazzocchi *et al.*, 2006). Angulo and Gil (2007) report consumption drops of 40% in France, 60% in Germany, 42% in Italy and 30% in Portugal in the first months after the crisis. These studies typically model AIDS demand systems of beef and other meats, and include some index of media coverage to take into account the effect of information. They are based on statistical aggregate data, since in general little information is available on consumers' individual immediate reactions (one exception is Smith *et al.*, 1999) and on later individual behaviour. The estimates of demand systems suggest that short-term reaction is very strong, but long-term effects are much weaker. The extent of the recovery is different between studies and between countries. Burton *et al.* estimate that the long-run effect of the first wave of BSE is a 4.9% decrease of the expenditure share in beef in UK; the structural reduction of the expenditure share evaluated for the Netherlands by Mangen and Burrell is 2.5%; and Mazzocchi *et al.* evaluate a 3.3% structural decrease of the expenditure share of beef in Italy. Due to the nature of the AIDS model, estimates are in terms of expenditure shares, and the overall quantitative effect remains more uncertain. What is nevertheless clear from these studies is that a typical consumers' reaction to scares concerning beef is shifting to consumption of other types of meat.

Studies relying on aggregated data can obviously only observe consumers' choices among existing meats. But governments and producers also can, and actually did react to the scares trying to provide assurance of food safety through certification or traceability (see Sans *et al*. (2008) for a discussion of the French case), thus creating qualitatively different products, that might be considered safer by consumers. Finding methods for assessing *a priori* which would be the reactions to alternative products is therefore of utmost importance. Accordingly, another stream of literature tries to ascertain consumers' attitudes towards safer or certified products not yet actually available. Certification may concern the origin, when the danger is felt by consumers as coming from foreign products, or traceability. Another obvious alternative to regular beef in terms of safety is organic beef, that must be produced in "natural" ways and is subject to control. Unlike the previously cited studies, the stream of literature dealing with certified products relies on direct surveys of consumers rather than on aggregate data and is based on stated preference methods, since these products were not available to consumers. Examples of this stream are: Latouche *et al.* (1999), who investigate French consumers' willingness-to-pay (WTP) for an hypothetical beef not transmitting BSE; McCluskey *et al.* (2005), on WTP for BSE-tested beef in Japan; Angulo and Gil (2007) who model consumers' WTP for certified beef in Spain. The estimated premium for safer beef range from 5% in Spain (Angulo and Gil) and 5-7% in France (Latouche *et al.*) to more than 50% in Japan (McCluskey *et al.*). The setting of these papers is similar to the one widely adopted to assess consumers' WTP for some higher quality product: consumers are asked to state their willingness-to-pay a price, or a premium relative to the regular good, using open-ended or single- or double-bounded elicitation formats (*e.g.*, Mullen and Wohlgenant, 1991; Henson, 1996; Fu *et al.*, 1999; Boland *et al.*, 1999; Loureiro and Hine, 2002; Gil *et al.*, 2000, among several others) [1]. Unfortunately, such an approach, though very frequent in the literature, is questionable for goods whose quantity can be freely chosen by consumers (Corsi, 2007). The WTP for a particular quality, if expressed in terms of price, is the *marginal* WTP and, hence, depends on the chosen quantity. If the quantity is not indicated when asking the price the consumer is willing to pay, or the premium relative to a regular quality, the question is ambiguous and the results depend on which interpretation of the question respondents imply, whether the price they would pay for the quantity they have in mind or the price for which the purchased quantity falls to zero. Things are even worse when the new quality product does not completely substitute for the old quality one, since consumers may also decide to consume both the old and the new product.

The goal of this paper is therefore estimating consumers' WTP for organic beef at a short and longer term after the BSE crisis explicitly keeping into account the possibility for consumers to adapt the purchased quantity of the new product. This is done using an innovative stated preference method. Since the paper is based on two surveys, the first one conducted in June and July 2001 (few months after the BSE crisis) and the second one in April and June 2003, it also contributes to understand consumers' short and long-term reactions to the "mad cow" crisis.

In the following paragraph, the theoretical approach to estimation of consumers' willingness to pay in terms of price is presented. Paragraph 3 presents the econometric strategy used for estimation. In paragraph 4, a short description of data used is presented. To put consumers' WTP for organic beef in perspective in the two periods, paragraph 5 analyses consumers' stated reactions to the BSE crisis in terms of regular beef consumption. In paragraph 5, their willingness-to-pay for organic beef is estimated, and the comparison between 2001 and 2003 results allows for an assessment of the long-term impact of BSE on willingness-to-pay for organic beef. Some conclusions follow.

[1] Alternatively, choice experiments are used for assessing consumers' WTP for some characteristic of the product (*e.g.*, Carlsson *et al.*, 2007; Lusk *et al.*, 2003; among others).

2. Theoretical approach

As already noted, if willingness to pay is expressed in terms of price, it is contingent on the chosen quantity. Hence, maximum WTP for a good in terms of its price means the maximum price consumers are willing to pay for it, *i.e.*, its reservation price. The theoretical background for assessing WTP for a new product like organic beef (not yet commonly available at the time of the surveys) is as follows. If a consumer makes his/her choice when organic beef is not available choosing the optimal quantity q_0^0 of regular beef at price p_0 and achieving utility v_0, the expenditure function, indicating the minimum expenditure needed to achieve utility v_0, is:

$$e_0 = e_0(\mathbf{P}, p_0, v_0) = e_0(\mathbf{P}, p_0, v(\mathbf{P}, p_0, \mathbf{s}, M)) = e_0(\mathbf{P}, p_0, \mathbf{s}, M) \quad (1)$$

where $\mathbf{P}$ is the vector of other prices, $\mathbf{s}$ are preference shifters such as attributes of the individual, and M is income.

When organic beef becomes available at price p_1, to attain the same utility level v_0 the minimum expenditure will be:

$$e_1 = e_1(\mathbf{P}, p_0, p_1, v_0) = e_1(\mathbf{P}, p_0, p_1, v_0(\mathbf{P}, p_0, p_1, \mathbf{s}, M)) = e_1(\mathbf{P}, p_0, p_1, \mathbf{s}, M) \quad (2)$$

where price p_0 is included in the expenditure function because regular beef is still available.

The consumer will purchase some organic beef if the expenditure now needed for reaching the same utility is less, i.e., if:

$$e_1(\mathbf{P}, p_0, p_1, s, M) < e_0(\mathbf{P}, p_0, \mathbf{s}, M) \quad (3)$$

or:

$$d(\mathbf{P}, p_0, p_1, \mathbf{s}, M) > 0 \quad (4)$$

where $d(\cdot) = e_0(\mathbf{P}, p_0, \mathbf{s}, M) - e_1(\mathbf{P}, p_0, p_1, \mathbf{s}, M)$ can be called the difference-in-expenditure (DE) function.

The DE function is decreasing in p_1, since e_1 is increasing in p_1 and p_1 is not an argument in e_0. For a given price p_1^* the DE reduces to zero and, for any $p_1 > p_1^*$, the difference in expenditure remains zero: the consumer would simply buy the same quantity of regular beef, and no organic beef. Hence, p_1^*, the reservation price, is the maximum price consumers are willing to pay for organic beef.

3. Econometric strategy

If the DE is expressed as a function of explanatory variables and of a random term, then assuming an appropriate distribution for the random term allows estimation of the DE function with maximum likelihood methods. For an empirical analysis of the problem, following the random utility model, it is assumed that, while consumers know their preferences with certainty, there are some components unknown to the researcher that are treated as random. Calling ε_0 and ε_1 the random components, and e^s_0 and e^s_1 the systematic components of expenditure functions (1) and (2), respectively, the condition for the consumer to buy a positive quantity of organic beef is:

$$e^s_1(\mathbf{P}, p_0, p_1, \mathbf{s}, \mathrm{M}) + \varepsilon_1 < e^s_0(\mathbf{P}, p_0, \mathbf{s}, \mathrm{M}) + \varepsilon_0 \quad (5)$$

or:

$$d^s(\mathbf{P}, p_0, p_1, \mathbf{s}, M) > \mu, \tag{6}$$

where $d^s(\cdot) = e^s{}_0(\mathbf{P}, p_0, \mathbf{s}, M) - e^s{}_1(\mathbf{P}, p_0, p_1, \mathbf{s}, M)$; $\mu = \varepsilon_1 - \varepsilon_0$; and $d(\cdot) = d^s(\cdot) - \mu$.

To estimate the DE equation, a density function has to be assumed for μ. Since $d(\cdot)\geq 0$, then $\mu < d^s(p_1^*)$ when the consumer chooses some organic beef, and $\mu = d^s(p_1^*)$ otherwise. Hence, the density function of μ must have a mass density at $d^s(p_1^*)$. Therefore, in our exercise μ is assumed to have a normal probability distribution, censored at $d(p_1^*)$. It is then possible to express the probability of a positive consumption of organic beef for a particular p_1 offered (p_{bid}) in terms of the cumulative density function of μ, G_μ, provided that $p_{bid} < p_1$[2]. The probability that a consumer will respond "yes" to an offered p_{bid} is the probability that μ is smaller than $d^s(\cdot)$ or:

$$P(\text{consumption}) = P[\mu < d^s(\mathbf{P}, p_0, p_{bid}, \mathbf{s}, M)] = G_\mu[d^s(\mathbf{P}, p_0, p_{bid}, \mathbf{s}, M)] \tag{7}$$

and:

$$P(\text{no consumption}) = 1 - G_\mu[\cdot]. \tag{8}$$

Maximum likelihood techniques can be employed to estimate the parameters in $d^s(\cdot)$. It is important to note that with this approach, if the consumer is willing to buy some organic beef, even a lower quantity than the quantity of regular beef he/she used to buy before organic beef was made available, this should be considered as a "yes" response [3]. Also observations from persons who presently do not consume regular beef can be used (in our sample, some consumers had stopped to consume beef, due to the BSE). They are presently at a corner solution for regular beef, but if organic beef is made available, they may decide to consume it, if their expenditure when organic beef is available is less than their expenditure when it was not, holding utility constant [4].

Explanatory variables comprise the prices of regular and organic beef, income, and taste shifters such as socio-economic characteristics (prices of other goods were assumed to be the same for all consumers and were therefore not included).

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[2] The condition that $p_{bid} < p_1^*$ has to be checked after the estimation. This is because the probability of a "yes" response for any $p_{bid} > p_1^*$ is $G_\mu(p_1^*)$ and not $G_\mu(p_{bid})$, and because p_1^* is unknown before the estimation. In practice for estimation one has to use the regular normal cdf, that gives the exact probability of a "yes" response for any $p_{bid} < p_1^*$. If $p_{bid} > p_1^*$, the relevant probability is underestimated.

[3] Our approach is similar to Cameron's (1988) treatment of referendum contingent valuation questions in that it uses the difference in expenditure (see also Hanemann and Kanninen, 1999). Nevertheless, in Cameron's approach the difference in expenditure measures the willingness to pay for a *given* change in the quantity/quality of the relevant good; put in the same terms, in our approach it measures the willingness to pay for an *unknown* (to the researcher) quantity of the new good at a given price, *allowing for a change in the quantity of the regular one.*

[4] It should be noted that, since the DE function is a difference between two expenditure functions, it is quite possible that income and personal characteristics effects vanish in the DE equation if their coefficients were equal in both expenditure equations. Nevertheless, in our exercise we preferred to keep them, in order to take into account possible changes in the coefficients when organic beef becomes available.

Since the reservation price p_1^* is the price for which the expenditure functions with and without organic beef are equal, *i.e.* the level of p_1 for which the difference in expenditure is equal to zero, p_1^* can be recovered by setting $d(\cdot)$ to zero and solving for p_1, which yields a reservation price (RP) equation, *i.e.,* an equation giving the maximum price consumers are willing to pay as a function of the explanatory variables other than p_1 [5].

In a random sampling, the distribution of the covariates is an unbiased estimator of their distribution in the reference population. The reservation price for each consumer in the sample can then be calculated multiplying the individual covariates by the vector of the coefficients of the RP equation, and the sample mean and other descriptive statistics for the sample can be used to estimate the relevant parameters in the population. Confidence intervals for the estimates can also be found by simulation methods (Krinsky and Robb, 1986). Multiple random drawings from a multivariate normal distribution with mean β (the vector of the estimates of the DE equation) and variance-covariance matrix V (the estimated variance-covariance matrix) result in random β vectors; from each of them, a new vector of the RP equation coefficients can be calculated, and the reservation prices for the sample can be computed. The final results are empirical distributions of the average reservation prices. From these, $(1-\alpha)$ confidence intervals are obtained by sorting the distributions and dropping $\alpha/2$ values from both tails of the sorted distributions.

4. Data

Till European Council Regulation (EC) 1804/1999 was issued, no animal product in Europe had the right to be labelled as "organic", but since a national regulation was further needed, in Italy it was not before 2000 that organic animal products could be legally marketed, so that for most Italian consumers organic beef was not actually available at the time of the first survey. Actually, due to technical production problems connected with the regulations and to market uncertainties, the production of organic beef in Italy is still very rare nowadays, and it was more so at the time of the surveys. Investigating consumers' attitudes and willingness-to-pay for organic beef was an important goal of the surveys. The first research was indeed promoted by a regional organic farmers' association (Agri.Bio Piemonte), that was interested in market prospects of organic beef. Since organic beef was not commonly available for consumers, a stated preference technique was needed. Hence, a questionnaire was prepared to this purpose.

Organic beef is an obvious alternative to regular beef in times of food scare, since in principle it should not suffer from the same concerns as regular beef, because organic stock-raising is based on pasture and on organic animal feed, which should prevent any BSE problem. It was therefore easily predictable that the weakening of the scare would affect consumers' attitudes towards organic beef, which was of empirical and theoretical interest. This was the reason for the second survey, based on our own University funding.

[5] While, as in usual probit and logit analysis, the parameters in the DE equation are only identifiable up to a scale parameter, if the DE equation is linear in p_{bid} the parameters of the RP equation are perfectly identified, since they are found by dividing the coefficients of the difference-in-expenditure equation other than the p_{bid} by (-) the coefficient of the p_{bid}.

Data for both surveys were collected through random telephone surveys in Piedmont (Italy). The target population was those residents in Piedmont Region usually in charge of buying food for themselves and their family. The interviewers therefore explicitly asked to talk with the household member usually in charge of buying food. The response rates were 51.4 and 57.7 percent, respectively in 2001 and 2003, which is reasonably high for a telephone survey. The interviewers stopped some interviews when respondents were found to be permanently out of the beef market (vegetarians, people consuming only other meat for health reasons, farmers self-consuming their products). Finally, after eliminating questionnaires that were not usable because they were incomplete, the 2001 survey consisted of 400 valid questionnaires, and the 2003 survey of 326. The first survey was designed with the goals of: i) analyzing consumers' familiarity with, and purchase habits of, organic products, ii) evaluating consumers' willingness to pay for organic beef, iii) determining consumers' preferences about organic beef selling outlets, packaging and label; also their reactions to the BSE crisis were investigated. More details on this survey can be found in Corsi and Novelli (2007). The second survey skipped questions about preferences for organic outlets and packaging, but in addition investigated whether consumers had changed their first reactions to the BSE or the BSE had induced long-term changes in consumption behaviour. Table 1 presents the descriptive statistics of the socio-economic characteristics of the survey and of the prices of regular beef cuts as reported by the respondents.

Table 1. **Descriptive statistics of the variables used for estimation**

	2001			2003		
	Mean	**Std. dev.**	**Obs**	**Mean**	**Std. dev.**	**Obs**
Price of roast *(Euro/kg)*	13.234	2.542	200	12.587	3.956	109
Price of minute steak *(Euro/kg)*	15.260	2.888	226	14.378	3.823	127
Age	50.108	15.612	400	53.218	15.715	326
Education *(years of study)*	10.340	3.843	400	9.298	4.008	326
Household size	3.188	1.054	400	2.687	1.104	326
Big city *(= 1 if living in towns with more than 50,000 inhabitants)*	0.310	0.463	400	0.313	0.464	326
Familiar with organic products *(1 = yes)*	0.640	0.481	400	0.724	0.448	326
Sex *(female = 1)*	0.818	0.387	400	0.813	0.391	326
Income class 1	0.077	0.267	400	0.117	0.321	326
Income class 2	0.308	0.462	400	0.285	0.452	326
Income class 3	0.340	0.474	400	0.209	0.407	326
Income class 4	0.195	0.397	400	0.104	0.306	326
Income class 5	0.080	0.272	400	0.285	0.452	326

Notes:

1. Income classes are as follows: 1 = 0-7,747 €/year; 2 = 7,747-15,494 €/year; 3 = 15,494-23,241 €/year; 4 = 23,241-30,987 €/year; 5 = over 30,987 €/year (Lira values converted to Euro).

2. The table excludes those observations with missing data for variables used in the estimations (the full samples were 402 observations in 2001 and 330 in 2003). Missing data about prices are referred to those people who could not remember the price they paid for regular beef at the moment of the interview.

Some significant differences can be identified in comparing the two surveys. The mean age is significantly higher in the second one (as measured with a test of the difference in means), while the average household size is smaller, thus reflecting the general aging of population and demographic changes. The price of minute steak is significantly lower in 2003 than in 2001, while for roast the difference is not statistically significant (these prices were directly asked to respondents who consumed the specific cut, which explains the differences in the number of observations). This suggests that the mad cow crisis depressed the beef marked also at long term. Familiarity with organic products is significantly higher in 2003, indicating the spread of organic consumption. The decline in average education, though significant, is not intuitive, and we are inclined to think it is an unusual random result.

The crucial point of the questionnaires concerned the price consumers were willing to pay for organic beef. Since WTP for a particular quality expressed in terms of price depends on the chosen quantity, in these surveys quantity adjustments to the introduction of organic beef made available at different prices were considered. Since evaluating a new product like organic beef and deciding how much of it one is prepared to consume at a specific price is a demanding task in cognitive terms, it was decided to use as a quantity reference the previous consumption of regular beef. For respondents, it is much easier to decide whether they would consume the same, more, or less of organic beef than they used to consume of regular beef, if any.

Therefore, in the interview, after an explanation about the prospective availability, the characteristics, and the certification process of organic beef, those respondents presently consuming regular beef were asked whether, if organic beef were made available at a specific price (bid price) they would consume it. Three answers were prompted: "Yes, I would buy it in the same quantity I'm currently consuming" [6]; "Yes, but I would buy less than what I'm currently consuming"; "No". Respondents who had given up eating beef after the "mad cow" events were asked about the possibility to go back and consume it; in this case, the answer could only be "yes" or "no". WTP for two beef cuts largely popular among Italian consumers, roast and minute steak, were evaluated; the former cut is cheaper but more time-consuming for cooking. Those respondents who did not like a specific cut were therefore excluded from the specific estimation.

To increase the elicitation process efficiency, a follow-up question was used (Carson *et al.*, 1986; Hanemann *et al.*, 1991): those respondents who had answered 'yes' to the first question were asked again if they were willing to pay a second and higher price; if the answer to the first question was 'no' the interviewer proposed a lower price. Three initial bids were chosen, the same for both surveys [7]. The first bid prices were

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[6] In retrospect, it would be more precise to word the first answer as "the same or a greater quantity", since in principle respondents might have consumed more organic beef than they did of regular beef. Nevertheless, no respondent objected that he/she would buy more organic beef, if available, than regular one. Moreover, in terms of econometric estimation (see below) this is immaterial, since the implied portions of the cdf used for estimation are the same for "same quantity" and for "more".

[7] Since between the first and the second survey Euro was introduced, in the second survey bid prices were given in Euro, but also the corresponding value in Italian Lira was proposed. In the following, the Lira values are converted to Euro.

set for roast at 25,000, 30,000, and 35,000 Italian Lire/kg in the 2001 survey, and to the corresponding 13.50, 15.50, and 18.00 Euro/kg in 2003 (the conversion to Euro of national currencies took place between the two surveys). The values for minute steak were 30,000, 35,000 and 40,000 ITL/kg, and 15.50, 18.00 and 20.60 Euro/kg in 2001 and 2003, respectively. Bid prices were set at levels considered a priori higher than, or equal to, first-rate quality beef currently on sale. In fact, average prices for regular roast were 13.23 and 12.59 Euro/kg in 2001 and 2003, respectively, and 15.26 and 14.38 Euro/kg for minute steak (see Table 1), though these prices show a large variation. Market actual prices are lower in 2003 than in 2001, a sign that the beef market still suffered in 2003 from the "mad cow" events of 2001. The second bid price was 5,000 ITL/kg (2.58 Euro/kg) higher or lower than the first bid price, depending on whether the respondent was willing or unwilling to pay the first price.

To avoid question order bias, six different versions of the questionnaire were randomly submitted to the respondents, each with different ordering of questions and/ or of prompted answers. The questionnaire was pre-tested with a small pilot sample in order to assess the adequacy of the bid design and the clearness of the questionnaire.

5. Reactions to the BSE crisis in terms of beef consumption

To put in perspective consumers' WTP for organic beef in the two periods, their reactions to the "mad cow" crisis is relevant. This can be assessed because the questionnaires asked about their consumption habits. In particular, respondents were asked whether they presently consumed beef and, if not, why. Those who stated they stopped eating beef because of the BSE were 9.7% and 2.1% of total consumers (excluding those who did not consume beef because they did not like it or because they were vegetarian, and whose interviews were stopped [8]) in 2001 and 2003, respectively, a statistically significant difference. The difference (7.6 points) can therefore be considered as the transitory "radical" change in beef consumption, and the remaining 2.1 percent the long-term one.

Nevertheless, changes in habits do not confine themselves in stopping beef consumption. Both in the first and in the second survey respondents were asked if they had changed their consumption habits at the times of the BSE scare and, if so, in which ways. In 2001 39.3 percent declared they had changed their habits, and in 2003 the corresponding figure was 36.1 percent (the difference is not statistically significant). A probit analysis of the determinants of consumption change [9] suggests that almost no socio-economic characteristic influenced the choice: the model is hardly significant for 2001, and not significant for 2003. Though some variables were

8 Respondents who answered they did not consume beef were asked the reason and were prompted three answers: 1) We don't eat it because we don't like it/we are vegetarian; 2) We stopped consuming because of the "mad cow" events; 3) other reasons (to be specified). No one stated the budget constraint as the "other reason", which is reasonable in a developed region like Piedmont. We thank a referee for asking to clarify this point.

9 For the sake of brevity, the results are not presented here, but are available from the authors upon request.

significant either in 2001 or in 2003, the unavoidable conclusion seems to be that the panic hit the overall population regardless of income and other socio-economic characteristics and mainly depending on the individual psychological impact.

Nevertheless, there are important differences in what respondents state in either year about the modalities of change in consumption at the times of the scare. The reactions could be: 1) giving up completely beef; 2) just reducing beef consumption; 3) reducing beef consumption but increasing other kinds of meat (chicken, pork) or fish; 4) stopping consuming certain beef cuts such as *e.g.* t-bone steaks (that could be considered as "dangerous"). A comparison between the answers in the two surveys is interesting to analyse how consumers recall their reactions. In both years (Table 2), the highest share of responses was "I consume/d less of beef and more of other kinds of meat (chicken, pork) or fish", but more so in 2001 than in 2003 (68.4 vs. 49.6 percent; a chi-square test shows the difference is significant). The share of those declaring they had given up eating beef is also declining (24.7 and 21 percent), but the difference is not significant. By contrast, in 2003 the shares of those stating they had simply reduced beef consumption (11.8 percent, as compared to 3.8 percent in 2001) and of those declaring they had stopped consuming certain cuts such as *e.g.* t-bone steaks (17.6 percent, as compared to 3.2 percent in 2001) are much higher. These differences between the two surveys (which chi-squared tests indicate as significant) are hardly explainable with the available information, but in a way it seems that while consumers correctly recall when they totally stopped consuming beef, they ex post rationalise their other choices, in particular when they ex post state they choose the most "rational" response (avoiding "dangerous" cuts).

Table 2. **Stated beef consumption changes immediately following BSE**

	2001	2003	2001	2003	2001	2003
	N.	N.	% over changed	% over changed	% over total	% over total
Gave up beef in 2001	39	25	24.7	21.0	9.7	7.6
of which: still do not consume beef in 2003		*7*		*5.9*		*2.1*
Less beef	6	14	3.8	11.8	1.5	4.2
Less beef, more of other meat/fish	108	59	68.4	49.6	26.9	17.9
Gave up some beef cuts	5	21	3.2	17.6	1.2	6.4
Total changed consumption habits	158	119	100.0	100.0	39.3	36.1
Total questionnaires	402	330			100.0	100.0

An interesting information concerns how consumers in 2003 stated their long-term reaction to the mad cow crisis. In the 2003 survey, respondent who stated they had changed their consumption habits because of the "mad cow" were asked how their consumption had changed since then. Of those who had changed their beef consumption habits after the BSE crisis, 55.9 percent had gone back to consuming the same quantity of beef as before the crisis. By contrast, about 29 percent of respondents stated they had maintained the changes, and 15 percent had increased their consumption as compared to the immediate period after the crisis, but still consumed less beef than before the crisis. This supports the view that the BSE crisis, though recovered to a large extent, left some permanent sign on consumers' behaviour. Again, there seems to be no particular socio-economic determinant of the decision to maintain the changes taken on in 2001. A probit model analysing the choice to keep to the changes adopted in 2001 is overall significant only at the 10 percent level, and the only individual significant coefficient is education [10]. Again, reactions to the BSE, the long-term ones in this case, seem to be rather random.

6. BSE crisis and willingness-to-pay for organic beef

The core of the questionnaires was the question about prospected consumption of organic beef according to different prices. Table 3 presents the share of responses for each bid, referred to the initial bid for beef consumers [11]. In both years, the share of "no" responses increases with increasing prices, with the exception of the highest bid price for roast in 2003. The shares of "yes, same quantity" and of "yes, but less" responses respectively decrease and increase with price but for the highest minute steak price in 2003. When comparing the surveys, the shares of "no" and of "yes, but less" responses are higher in 2003 than in 2001 for respondents asked the lower initial bids; the opposite holds in general for respondents asked the highest initial bid. Chi-squared tests on the distribution for each initial bid price reject the hypothesis of no effect of the year for all initial bids but for the highest one for roast. Though the consumption patterns are not very clear from these raw data, it might be argued that the time distance from the BSE crisis decreased the willingness to pay for organic beef among those consumers who were willing to buy it, but paying low prices, while it increases among the high price segment. This hypothesis can be better verified through the proposed estimation of the DE and RP equations.

The estimations concern two sub-samples (Group A and Group B). Group A is formed by those respondents who consumed regular beef at the time of the surveys and reported its price, which is therefore included among the variables in the equations (regular beef price was asked to respondent, and exhibits a large variation). Some observations are from persons who did not consume regular beef at the time of the surveys due to the BSE. These observations can be used but, since they already excluded consumption of regular beef, price of regular beef does not enter in their DE equation (and is also unknown), so that their DE equation has to be estimated separately. We pooled with these respondents those who consumed beef, but could not remember the price they paid. Group B is therefore formed by those who did not consume regular beef plus those who did consume regular beef but could not remember its price, which is therefore not included among the variables in the equations.

[10] Results are available from the authors upon request.

[11] This table does not include respondents not consuming beef at the time of the surveys or not consuming the specific cut. Respondents not consuming beef at the time of the second survey were very few, which makes the comparison with the first one somewhat problematic.

Table 3. **Shares of responses to the WTP question**

	2001			2003		
Roast						
Bid Price (Euro/kg)	*13.00*	*15.50*	*18.00*	*13.00*	*15.50*	*18.00*
Yes, same quantity	74.6	63.4	42.7	57.1	41.2	37.8
Yes, but less	19.5	27.7	36.4	27.6	38.2	44.9
No	5.9	8.9	20.9	15.2	20.6	17.3
Total	100.0	100.0	100.0	100.0	100.0	100.0
N.	*118*	*112*	*110*	*105*	*102*	*98*
Minute steak						
Bid Price (Euro/kg)	*15.50*	*18.00*	*20.60*	*15.50*	*18.00*	*20.60*
Yes, same quantity	72.1	62.8	20.8	49.1	34.3	35.9
Yes, but less	23.0	26.5	45.6	36.8	45.4	38.8
No	4.9	10.6	33.6	14.2	20.4	25.2
Total	100.0	100.0	100.0	100.0	100.0	100.0
N.	*122*	*113*	*125*	*106*	*108*	*103*

Note: This table does not include respondents not consuming beef at the time of the surveys or not consuming the specific cut

The DE equations for both surveys are estimated with the maximum likelihood method exploiting the double bounded format. There has been much discussion in the literature about possibly different WTP distributions implied by the initial and follow-up bid or, more generally, about bias introduced by the double-bounded (DB) format (Alberini *et al.*, 1997; Cameron and Quiggin, 1994; Cooper *et al.*, 2002; Herriges and Shogren, 1996; Kanninen, 1995; O'Conor *et al.*, 1999). We tried to re-estimate the model including a yea-saying effect as suggested in Kanninen (1995), but the model failed to converge. The method suggested by Herriges and Shogren (1996) for controlling for anchoring is not applicable in our case, due to the polytomous rather than dichotomous responses among beef consumers (yes, yes but less, no). Nevertheless, we checked for bias in three ways. First, we exploited the fact that the second bids were in some cases the same as the first bids (*e.g.* those who responded "no" to a first bid of 30,000 ITL/kg were asked a second bid of 25,000 ITL/kg, which was the first bid asked to other respondents) to test the independence of response distribution from the bid sequence. The results of the chi square tests confronting the distributions of the responses to the first and to the second bid for the same prices are mixed: for 10 over 16 pair wise tests we were unable to reject the hypothesis of independence. Second, we confronted the parameters of the DE equations estimated with the DB format to those estimated with a single bounded (SB) format, *i.e.*, only using responses to the first bid. For all equations, the five percent confidence intervals of the parameters were overlapping for all variables. We also tested with a likelihood ratio test the restrictions of the SB parameters being equal to the parameters estimated with the DB, and vice versa. We were never able to reject the restrictions. Third, we estimated the reservation prices with the SB format. The five percent confidence intervals of the reservation prices calculated with SB and DB formats are always overlapping; the differences are around 5-6 percent for the total of observations and, more importantly, are of different signs for the different groups. We conclude that there is no strong evidence of a double-bounded format bias. These results are available from the authors upon request.

The DE parameters (Table 4) show how the explanatory variables influence the probability of a positive response, *i.e.*, the probability of consumption of any positive amount of organic beef. The coefficients of the bid price are always significant and, as predicted, negative. The coefficients of the price of regular beef in Group A are significant (though weakly significant for Roast in 2003) and positive, thus indicating that consumers paying more for regular beef are also more likely to purchase organic beef, ceteris paribus. Other variables that are significant and positive, though not for all cases, are familiarity with organic products and income. That consumption of organic products is higher in higher income population is well established. The fact that familiarity with organic products is a positive determinant of willingness to purchase suggests that the experience of consuming organic products is positive or, at least, that information about organic products has a positive effect on the willingness to purchase them. By contrast, other socio-economic variables like age, education, gender, and household size are never significant.

Unlike the raw data, there is weak evidence for an effect of the year. We tested with log-likelihood ratio tests the restriction of equal parameters in both years for the DE equations and were unable to reject the restriction (at the 5 percent level)[12]. Also using an alternative way of testing the year effect, namely inserting a year dummy in the equations estimated on the pooled sample, we found that the relevant parameters were never significant. Nevertheless, while the overall models are not significantly different, the parameters of the bid prices in the DE equations are larger in absolute terms in 2001 than in 2003, and their difference, as measured by a t-test, is significant. In other words, in 2003 an increase in the price would decrease the probability of a positive difference in expenditure and hence, the willingness to buy organic beef, less than what the same price increase would have done in 2001.

In Table 5 the relevant RP equations are shown. The parameters of the RP equations, *i.e.*, the equations giving the reservation prices as functions of the explanatory variables, are simply the coefficients of the DE equations divided by (–) the coefficient of the bid price, since we chose a linear specification for the DE equations. Confidence intervals are computed using Monte Carlo methods (3500 draws). The parameters that are significant are the same as in the DE equations. Just as examples of how to interpret the RP equations, the price coefficient in the RP equation for Group A in 2001 suggests that one Euro increase in the price the consumer pays for regular roast implies an increase of 0.966 €/kg in the maximum price he/she would pay for organic roast.

The average reservation prices are shown in Table 6. The reservation prices for the surveyed consumers have been estimated with the RP equations, and, using a Monte

12 Results for the pooled sample are not presented here to save space, but are available from the authors upon request.

Table 4. **Difference-in-expenditure equations**

	Roast				Minute steak			
	2001		2003		2001		2003	
	Coeff.	t-ratio	Coeff.	t-ratio	Coeff.	t-ratio	Coeff.	t-ratio
Group A								
Constant	– 0.387	– 0.295	0.668	0.325	2.139**	1.962	– 1.192	– 0.765
pbid	– 0.308***	– 4.851	– 0.172***	– 3.293	– 0.254***	– 6.891	– 0.112***	– 3.102
p	0.298***	4.471	0.154*	1.869	0.112***	2.659	0.143***	2.199
Age	0.016	1.436	0.003	0.161	0.013	1.591	0.015	1.011
Education (years)	0.027	0.513	0.091	1.164	– 0.011	– 0.284	0.074	1.144
Household size	– 0.032	– 0.229	– 0.130	– 0.598	– 0.014	– 0.145	0.064	0.342
Big town (1 ≥ 50,000 inh.)	0.568*	1.675	0.167	0.427	0.173	0.662	0.577	1.601
Familiar with organic	0.432	1.510	– 0.015	– 0.038	0.730***	3.068	0.223	0.646
Sex (Female = 1)	0.484	1.332	0.561	1.487	0.305	1.030	0.528	1.544
Income class 2	1.192***	2.654	– 0.036	– 0.058	0.656*	1.745	– 0.402	– 0.787
Income class 3	1.384***	2.666	0.475	0.677	0.897**	2.114	– 0.158	– 0.281
Income class 4	0.844	1.576	– 0.097	– 0.097	0.838*	1.843	– 0.019	– 0.021
Income class 5	0.738	0.984	0.010	0.016	0.692	1.248	– 0.201	– 0.373
N	199		109		226		127	
Log-likelihood	– 85.590		– 68.362		– 136.070		– 81.286	
Group B								
Constant	1.142	0.931	0.030	0.035	2.215**	2.213	0.839	0.810
pbid	– 0.116***	– 4.427	– 0.060***	– 3.315	– 0.164***	– 6.008	– 0.059***	– 3.306
Age	0.015	1.266	0.010	1.187	0.007	0.767	0.003	0.260
Education (years)	0.048	0.918	0.034	0.806	0.075	1.631	0.053	1.199
Household size	– 0.149	– 1.136	– 0.118	– 1.069	– 0.125	– 1.031	– 0.101	– 0.871
Big town (1 ≥ 50,000 inh.)	– 0.079	– 0.276	0.646*	1.875	– 0.354	– 1.447	0.279	0.902
Familiar with organic	0.527**	2.181	0.327	1.343	0.489**	2.234	0.133	0.525
Sex (Female = 1)	0.270	0.952	0.306	1.025	0.270	0.942	0.227	0.741
Income class 2	0.575	1.133	0.751**	1.984	0.610	1.588	0.479	1.333
Income class 3	0.773	1.228	0.872*	1.779	0.648	1.465	0.837**	1.997
Income class 4	0.110	0.185	1.302**	2.473	– 0.005	– 0.011	0.709	1.437
Income class 5	0.858	1.101	0.586	1.427	1.167*	1.653	0.320	0.843
N	177		197		171		192	
Log-likelihood	– 113.297		– 122.482		– 133.545		– 115.652	

***, **, * = significant at 1%, 5%, 10% level.

Table 5. **Reservation price equations**

	2001			2003		
	Coeff.	95% c.i.[a]		Coeff.	95% c.i.[a]	
		Lower b.	Upper b.		Lower b.	Upper b.
			Roast			
Group A						
Constant	– 1.256	– 12.581	6.337	3.893	– 35.987	24.113
p	0,966***	0.678	1.344	0,895*	– 0.023	2.681
Age	0.052	– 0.018	0.141	0.018	– 0.225	0.308
Education (years)	0.086	– 0.247	0.494	0.532	– 0.351	2.197
Household size	– 0.105	– 1.154	0.776	– 0.757	– 4.229	2.012
Big town (1 ≥ 50000 inh.)	1,842*	– 0.199	4.549	0.972	– 5.019	5.615
Familiar with organic	1.400	– 0.496	3.486	– 0.088	– 5.266	5.597
Sex (Female = 1)	1.570	– 0.766	4.566	3.267	– 1.017	11.608
Income class 2	3,867***	0.945	7.609	– 0.211	– 9.509	8.065
Income class 3	4,488**	1.277	8.687	2.771	– 6.267	13.922
Income class 4	2.739	– 0.745	7.137	– 0.564	– 15.399	12.785
Income class 5	2.392	– 2.599	7.891	0.056	– 8.546	8.712
Group B						
Constant	9.815	– 12.522	32.735	0.504	– 39.887	29.841
Age	0.126	– 0.080	0.355	0.172	– 0.125	0.650
Education (years)	0.415	– 0.567	1.415	0.563	– 0.901	2.711
Household size	– 1.283	– 4.021	0.998	– 1.961	– 7.981	1.851
Big town (1 ≥ 50000 inh.)	– 0.683	– 6.997	4.101	10,727*	– 0.546	26.809
Familiar with organic	4,527*	0.418	10.404	5.419	– 3.615	17.367
Sex (Female = 1)	2.319	– 3.042	7.626	5.075	– 6.346	16.967
Income class 2	4.940	– 3.931	14.318	12,455**	0.273	36.394
Income class 3	6.647	– 3.861	19.599	14,468*	– 1.080	41.720
Income class 4	0.943	– 9.802	11.865	21,601**	4.912	63.373
Income class 5	7.376	– 6.029	22.947	9.731	– 3.361	32.131
			Minute steak			
Group A						
Constant	8,413*	– 0.117	16.137	– 10.625	– 68.075	14.972
p	0,440**	0.130	0.745	1.274*	0.181	2.929
Age	0,052	– 0.011	0.127	0.136	– 0.141	0.589
Education (years)	– 0.042	– 0.356	0.248	0.660	– 0.478	2.812
Household size	– 0.054	– 0.795	0.687	0.568	– 3.476	5.012
Big town (1 ≥ 50000 inh.)	0.680	– 1.290	2.834	5.143	– 1.478	16.463
Familiar with organic	2,871***	0.903	5.257	1.988	– 5.196	9.622
Sex (Female = 1)	1.201182	– 1.123	3.575	4.708	– 1.093	18.091
Income class 2	2,580*	– 0.351	5.633	– 3.580	– 14.879	6.864
Income class 3	3,527**	0.295	7.342	– 1.411	– 13.989	11.223
Income class 4	3,296*	– 0.235	7.508	– 0.172	– 20.542	19.797
Income class 5	2.721	– 1.483	7.731	– 1.794	– 13.464	10.420

Table 5. **Reservation price equations *(continued)***

	2001			2003		
	Coeff.	95% c.i.[a]		Coeff.	95% c.i.[a]	
		Lower b.	Upper b.		Lower b.	Upper b.
Group B	13,485**	1.766	24.875	14.183	– 29.796	47.935
Constant	0.040	– 0.068	0.148	0.043	– 0.298	0.546
Age	0.454	– 0.101	1.098	0.904	– 0.573	3.675
Education (years)	– 0.759	– 2.372	0.701	– 1.713	– 7.749	2.423
Household size	– 2.157	– 5.737	0.683	4.713	– 7.287	16.869
Big town (1 ≥ 50000 inh.)	2,977**	0.366	6.147	2.254	– 7.819	13.611
Familiar with organic	1.643	– 1.841	5.495	3.829	– 8.284	17.077
Sex (Female = 1)	3.713	– 0.833	8.541	8.097	– 4.199	26.207
Income class 2	3.947	– 1.447	9.550	14.150**	0.678	44.859
Income class 3	– 0.032	– 6.309	5.440	11.983	– 4.484	41.224
Income class 4	7,104*	– 1.362	17.170	5.408	– 7.668	24.470
Income class 5	8,413*	– 0.117	16.137	– 10.625	– 68.075	14.972

*, **, *** indicate that the empirical 10%, 5%, 1% intervals do not include 0.

[a] Results from 3,500 Monte Carlo random draws.

Carlo simulation (10,000 draws), their mean, median and 95 percent confidence intervals have been computed for the sample[13]. Reservation prices are consistently higher in 2003 than in 2001. Confidence intervals (95 percent) in 2001 never include the 2003 average, for both Groups A and B for roast and minute steak. Confidence intervals in 2003 are much larger than in 2001, partly due to the smaller sample size, but also possibly to a larger dispersion in behaviour. Nevertheless, they almost never include the 2001 average, with the exception of Group A for roast. These results suggest that reservation prices did increase from 2001 to 2003.

This might seem at odds with the previous results that the shares of yes responses are lower at the proposed bid prices. Though, the discrepancy is only apparent. The parameters of the DE equations suggest that the demand is more inelastic in 2003 than in 2001. At the same time, the reservation price (which is the intercept of the demand curve) is higher in 2003. This suggests that the demand curve shifted to the left in 2003 relative to 2001, but became steeper, in a way illustrated by Graph 1. This could explain why at the proposed bids the shares of yes responses are lower in 2003. To further check this hypothesis, the percentage of consumers that, at the different bid price levels, were likely to purchase organic beef was estimated. The DE equations were used to compute the difference in expenditure for each particular

[13] Through these simulations the condition that $p_{bid} < p^*_1$ can also be checked (see footnote 2). For 2001, the estimated reservation prices are smaller than the first bid in a quite limited number of cases: 4.9 and 9 percent for Group A, and 8.8 and 2.3 percent for Group B, for minute steak and roast, respectively. For 2003, the relevant data are 5.5 and 5.5 percent for Group A, and 0.5 and 1 percent for Group B, for minute steak and roast, respectively. The shares are lower, and zero in most cases, if the comparison is made with the second bid. Therefore, it can be concluded that the violation of the condition is not a serious problem.

Table 6. **Estimated reservation prices (Euro/kg)** [a]

	2001					2003				
	Mean	Median	St. Dev.	95% c. i.		Mean	Median	St. Dev.	95% c. i.	
				Lower bound	Upper bound				Lower bound	Upper bound
Roast										
Group A	21.11	20.96	22.46	19.75	23.4	23.45	22.38	22.46	19.29	32.28
Group B	25.55	25.18	19.99	22.28	31.49	34.47	32.35	19.99	25.93	53.27
Total	23.33	22.74	21.97	19.92	30.13	28.96	27.39	21.97	19.68	47.95
Minute steak										
Group A	23.31	23.23	22.46	22.15	24.96	28.51	27.32	10.08	23.66	40.28
Group B	23.25	23.13	19.99	21.61	25.54	36.69	34.41	22.63	27.82	56.44
Total	23.28	23.19	21.97	21.81	25.25	32.60	30.71	17.99	24.06	51.64

[a] Results from 10,000 Monte Carlo random draws.

Graph. 1. **Organic beef demand curves in 2001 and 2003**

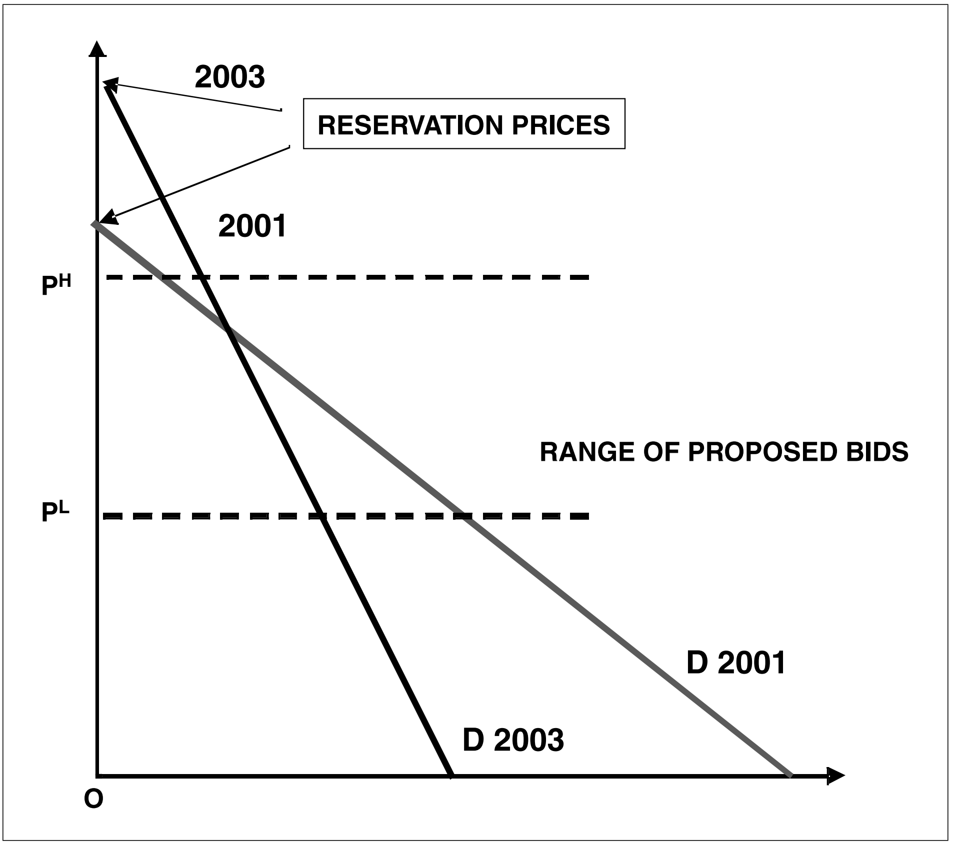

Table 7. **Estimated probabilities of purchase of organic beef at selected prices[a] and actual proportions of "yes" responses in the sample**

Price (Euro/kg)	Actual % of "yes" responses in the sample				Estimated average probabilities			
	Mean	95% C.I.	Mean	95% C.I.	Mean	95% C.I.	Mean	95% C.I.
	2001		2003		2001		2003	
				Roast				
Group A								
13.00	0.944	0.890-0.997	0.842	0.726-0.958	0.940	0.903-0.967	0.875	0.792-0.938
15.50	0.910	0.842-0.979	0.842	0.726-0.958	0.864	0.824-0.898	0.794	0.720-0.857
18.00	0.770	0.665-0.876	0.667	0.506-0.828	0.733	0.674-0.790	0.691	0.588-0.780
Group B								
13.00	0.919	0.852-0.987	0.843	0.758-0.928	0.891	0.837-0.935	0.837	0.782-0.884
15.50	0.895	0.815-0.974	0.750	0.644-0.856	0.837	0.783-0.883	0.804	0.751-0.851
18.00	0.724	0.609-0.839	0.921	0.854-0.987	0.765	0.702-0.822	0.767	0.708-0.819
				Minute steak				
Group A								
15.50	0.933	0.901-0.966	0.854	0.754-0.954	0.944	0.909-0.970	0.833	0.753-0.897
18.00	0.917	0.881-0.953	0.725	0.587-0.863	0.861	0.818-0.897	0.775	0.705-0.835
20.60	0.658	0.596-0.720	0.744	0.607-0.881	0.710	0.647-0.769	0.703	0.624-0.775
Group B								
15.50	0.917	0.875-0.958	0.862	0.773-0.951	0.858	0.798-0.909	0.838	0.783-0.885
18.00	0.839	0.784-0.894	0.817	0.727-0.907	0.762	0.703-0.817	0.804	0.751-0.853
20.60	0.582	0.508-0.656	0.730	0.621-0.840	0.635	0.562-0.705	0.764	0.701-0.821

[a] Results from 10,000 Monte Carlo random draws.

consumer in the sample at a particular bid price, multiplying the individual covariates by the vector of the coefficients of the DE equation. Then their probability to buy some organic beef was calculated [14]. Estimates of the proportion of likely buyers in the total population were then obtained by averaging the estimated probabilities of individuals in the sample, and the results were taken as an estimate of the probability

[14] Of course, this relies on the same assumption on the density distribution of the error as the one used in estimating the DE equation.

in the total population. Again, the Krinsky and Robb's (1986) simulation approach was used to provide empirical confidence intervals.

The results of this exercise are shown in Table 7, that also exhibits the actual response shares at the bid prices ("yes" responses in this table include both "yes, same quantity" and "yes, but less" responses). When comparing actual and simulated shares, it can be seen that the simulated mean values are sensibly similar to the actual ones, and that confidence intervals are in almost all cases overlapping. When comparing 2001 to 2003 values, the former are always higher for the lower bid prices, while in most cases they are lower for the highest bid prices. T-tests on the difference of the averages are always significant. This result, along with the result of higher reservation prices in 2003 supports the view of a steeper demand curve in 2003 than in 2001. In other words, these results suggest that the demand for organic beef was lower at the proposed bids in 2003 than it was in 2001, but that some consumers in 2003 were willing to purchase some organic beef up to prices that were higher than in 2001.

7. Conclusions

This paper had two goals: 1) analysing the short- and long-term reactions of consumers to the BSE crisis and 2) analysing if, and how, the time distance from the BSE crisis had affected consumers' willingness to buy organic beef.

Though the effect of the BSE crisis has weakened along with time distance, it left some permanent signs in consumers' behaviour. The share of those stopping eating beef, though significantly reduced from 2001 to 2003, was still positive. Probably more importantly, a sizeable part of those who had changed their beef consumption habits after the BSE crisis maintained, completely or partially, their changes. Both the choice of changing consumption habits just after the BSE crisis, and the choice of maintaining the change, appear to a large extent independent of socio-economic characteristics and hence seem to depend mostly on individual psychological characteristics.

The analysis of the effect of the time distance from the BSE crisis on consumers' attitudes towards organic beef leads to recognize that, at low prices, consumers were less willing to buy organic beef in 2003 than they were in 2001. Nevertheless, the share of those willing to buy it at high prices increased. The main conclusion is therefore that the demand for organic beef reduced, but that in the meantime it became more inelastic. From producers' perspective, these results suggest that organic beef was less likely to become a large consumption good in 2003 than it was in 2001, but that there were better prospects as a niche, high price product.

Finally, from the methodological point of view our results introduce a new methodology to evaluate consumers' response to an improved quality and show that, unlike the current literature, it is possible to take into account changes in consumed quantities when a different quality is prospected.

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