Consumer Welfare and the Loss Induced by Withheld Information: The Case of BSE in Italy

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Abstract

The paper develops a measure of consumer welfare losses associated with withheld information about BSE linkage with vCJD. food safety. The Cost of Ignorance (COI) is measured by comparing the utility of the informed choice with the utility of the uninformed one, under condition of improved information. Unlike previous work, based on a single equation demand model, the measure is obtained retrieving a cost function from a dynamic Almost Ideal Demand System. The results indicate that Italian consumers bore a significant loss because of the delayed release of information.

Keywords: Food safety, welfare analysis, information, BSE

JEL Classification: D80, D60, D12

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1. Introduction

Every food scare implies a welfare loss for households and a related reallocation of expenditure among goods because of the objective change in the quality (safety) level of available food. However, when relevant information about health threat is withheld for some time, a specific cost is borne by consumers, since they are not able to change their behaviour as they would have done if warned in due time. This cost has been termed by Foster and Just (1989) "cost of ignorance" (COI) and it takes the form of regret and sorry when consumers eventually are informed about the food safety problem and look back to the actions they could have undertaken (but they did not) to adapt to the new situation. The concern about the consequences of consuming unsafe food in ignorance strongly affects consumer decisions, especially when uncertainty persists after the release of information. The analysis of actual purchasing behaviour does not allow to give a complete assessment of the impact of new information. Thus, the monetary assessment through the COI approach has great relevance for evaluating the welfare effects of policy actions.

It is a well know feature of the BSE crisis that information about health risks has been undisclosed for some time. Drawing on the work by Foster and Just (1989), this study develops an approach to measurement of consumer welfare loss from the BSE crisis focusing on the effects of imperfect information.

The news about a possible linkage between BSE and its human counterpart (vCJD) in March 1996 triggered the most relevant food scare in Europe during the last years. The sudden fall in beef consumption all over Europe was just the first sign of a crisis which became structural under many respects, influencing the purchasing and consumption habits towards meats and food in general. A crucial issue in the BSE debate were the effects of withholding information about the potential risk for beef consumer. More specifically, the BSE shock highlighted the asymmetries in food safety information and undermined consumer trust towards institutions. We suggest to exploit the Cost of Ignorance measure to complete the economic assessment of the impact of the crisis by evaluating the impact on consumer welfare imputable to the concern for being unable to adjust his behaviour due to retained information.

The COI measure and the underlying concepts are described in section 2. A methodology for deriving a COI measure based upon the estimation of a dynamic Almost Ideal Demand System is presented in section 3, whereas the results an empirical implementation to the 1996 BSE scare in Italy are reported in section 4. The main findings of this study are summarised in section 5.

2. Cost of Ignorance: theoretical aspects

Within the traditional consumer surplus framework, welfare change measures are given by areas under the relevant demand curves. A paradoxical result may be obtained when new information reveals the potential hazards linked to the consumption of a particular good. As a consequence of this informational change, the demand curve for the affected good is likely to shift leftwards. Thus, using the area under the curve as a measure of consumer surplus leads to the incoherent conclusion that the provision of information makes consumers worse off. In other words, ignorance is blissful and no cost appear to be attached to it (Teisl and Roe, 1998).

To tackle the paradox we need to bear in mind the timing of the adaptation process triggered by the food crisis. Figure 1 illustrates the different steps involved in the decision process once the crisis has occurred. As information is withheld for some time, steps 0 and 1 take place. In this case, choices are made under ignorance and agents rely only on their prior beliefs, that is they assume food is still safe as it used to be. As information about the health threat is released, agents update their beliefs and a modified consumption pattern emerges.

Figure 1. Timing.

Step 0	Step 1	Step 2	Step 3	Step 4
		Time →		
Agent has prior beliefs about safety of food	Choices are made under ignorance resulting in allocation of consumption across goods	An information message is released	Agent updates beliefs via bayesan rule	Informed choices are made based on updated beliefs. A new allocation of consumption across goods takes place

The paradox of blissful ignorance arises because welfare comparison is made under different informational states. If the utility function (or alternatively the cost function) depends on the subjective perceptions of the safety level of food, then consumption under ignorance would yield higher utility than informed consumption. However, welfare measures should be undertaken under improved information because only in this state the different allocations of consumption (under ignorance and informed) can be correctly judged. Indeed, Cost of Ignorance (COI) is measured by comparing the utility of the informed choice with the utility of the uninformed one, under condition of improved information. This also explains why an ex post measure with respect to information release is needed In addition, COI refers to a given informational state that impinges on the subjective beliefs about the safety level of food. As the distribution is a subjective one, it can evolve in the course of time even if the underlying objective distribution no longer changes. This is a well known pattern of food scares, when huge press coverage of health threats causes an initial dramatic drop of consumption of the allegedly unsafe food (Burton, Young and Cromb, 1999; Verbeke and Ward, 2001). Initial panic is then followed by a slow, albeit often incomplete, recover of the previous level of consumption even after new information is released assuring that the crisis is over or that the actual level of risk is lower than supposed.

The delay in demand recovery may be explained by the asymmetric impact of good and bad news on consumption as it has been found by Liu (1998). This in turn relates to the role of trust in information sources. As it is well know, trust is fragile and once it is lost it may take a long time to recover its previous level. This is due to a number of reasons. First, negative events are more visible than positive ones. Second, negative events have a larger impact than positive ones because of their low frequency–high consequence nature. Third, sources of bad news tend to be trusted more than sources of good news. In addition, once distrust has arisen subsequent events tend to be interpreted in a distorted way leading to a reinforcement of previous beliefs (Slovich, 1993).

All these factors contribute to explain the observed evolution of beliefs and consumption behaviour in the course of time. As a consequence, different COI measures can be obtained by referring the informed status to ensuing dates.

Following this approach, a food safety crisis (as well as other quality change in goods or environment) may be represented by some shift in a quality parameter theta (θ) concerning a good, whose level of consumption (in quantity terms) is x.

$$\theta_0 \ge \theta_1 \qquad (1)$$

More generally, we may refer to θ as a vector of distribution parameters (such as mean and standard deviation) associated with uncertain quality. As illustrated above, it is the subjective estimate of θ rather than the objective value that is relevant, since only the former enters the consumer's utility or cost functions.

As the safety level of food is stochastic, consumers solve an expect utility maximisation problem whose solution can be represented by

$$x^* = f(p, y, \boldsymbol{\theta}) \qquad (2)$$

where x^* is the quantity of the good of interest which maximises the utility function, p is the price of the good of interest, and y is income, while $g=y-px^*$ is the expenditure on all other goods and is considered as a numeraire. Let the cost function, or the minimum expenditure attaining the level of utility U be:

$$e(p, \boldsymbol{\theta}, U) = \{\min(px + g); E_\theta[u(x, g, \boldsymbol{\theta}) \geq U]\} = y \qquad (3)$$

where q represents the (unknown) specific health effects associated to the consumption of the good.
Taking the absence of contamination as a baseline, the measure of consumer loss, when contamination occurs and consumers are informed without delay, is given by the compensating variation:

$$CV = e(p_0, \boldsymbol{\theta}_1, U_0) - e(p_0, \boldsymbol{\theta}_0, U_0) \qquad (4)$$

Where θ_1 now describes posterior beliefs that are the result of updating priors beliefs ($\boldsymbol{\theta_0}$) after the release of information about the food safety crisis. As information is promptly released, consumers are free to adapt their consumption bundles. The more likely reaction to the news is to reduce the consumption of the unsafe food so that $x_0^* = f(p_0, y, \boldsymbol{\theta}_0)$ will be greater than $x_1^* = f(p_0, y, \boldsymbol{\theta}_1)$.
Conversely, if consumers were not informed about the crisis, they would maintain the same consumption patterns that they adopted in the safe context. However, ex post, they would realise that a deterioration of the food safety levels has occurred and they would suffer a greater welfare loss, since they have not put in place any countermeasure. In this case, the measure for the correspondent welfare loss is the compensating surplus:

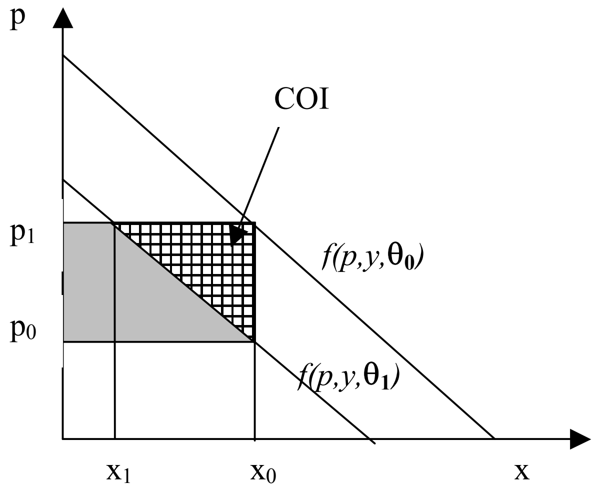
$$CS = e(p_0, \boldsymbol{\theta}_1, U_0 \,|\, x = x_0) - e(p_0, \boldsymbol{\theta}_0, U_0) \qquad (5)$$

The first term on the right hand side of equation (5) is a restricted cost function, as consumer choice is restricted to the bundle chosen under the original set of beliefs. In other words, consumers are forced to consume the same quantities that they would have chosen under ignorance even if they are actually aware of the health threat. Thus, an excessive quantity of the possibly unsafe food is consumed. The second term on the right hand side is a standard cost function, but it is measured before the release of information.
Since CS and CV differ only because of the restricted choice in CS, the difference between the two measure provides the cost related to ignorance:

$$COI = CS - CV = e(p_0, \boldsymbol{\theta}_1, U_0 \,|\, x = x_0) - e(p_0, \boldsymbol{\theta}_1, U_0) \qquad (6)$$

Noticeably, both terms of the COI expression refer to the same subjective distribution of quality parameters (θ_1). That is, consumer welfare is measured under condition of improved information by comparing the consequence of informed actions (consumption of quantity x_1) with those of uninformed actions (consumption of quantity x_0). In this situation, COI is a measure of the welfare improvement following the adoption of self-protection activities such as avoiding consumption of unhealthy food and self-insurance activities aimed to reduce the prospective severity of the health consequences.

Figure 2. Cost of ignorance



Since constrained cost function are difficult to retrieve, Foster and Just (1989) suggest an alternative strategy to measure COI:

$$COI = e(p_1, \boldsymbol{\theta}_1, U_0) - e(p_0, \boldsymbol{\theta}_1, U_0) + (p_0 - p_1)x_0 \qquad (7)$$

where p_1 is defined as the price "[…] that would need to be charged to cause the individual to choose x_0 given a Hicksian demand curve conditioned on U_0 and $\boldsymbol{\theta}_1$" (Teisl et al., 2001)[1].The measure is illustrated in Figure 2 as the difference between the area $(p_0-p_1)x_0$ and the area under the informed Hicksian demand curve between the two price lines corresponding to p_1 and p_0.
The measure can be adapted to deal with changes in quality of n goods. Defining $\mathbf{p_1}$ as the $n\times 1$ vector of prices such that the compensated demand with the perceived quality level $\boldsymbol{\theta}_1$ is equal to the initial level of consumption, represented by the $n\times 1$ vector $\mathbf{x_0}$. The quality levels before and after the disclosure of information are now represented by the $n\times m$ matrices $\boldsymbol{\Theta}_0$ and $\boldsymbol{\Theta}_1$, containing the m distribution parameters associated with each good. It is straightforward to derive the cost function as $e(\mathbf{p_0}, U_0, \boldsymbol{\Theta})$ and the COI expression is then given by:

$$COI = e(\mathbf{p_1}, U_0, \boldsymbol{\Theta}) - e(\mathbf{p_0}, U_0, \boldsymbol{\Theta}) + (\mathbf{p_0} - \mathbf{p_1})' \mathbf{x_0} \qquad (8)$$

It is worth noticing that while the adopted framework allows for changes in quality of all goods, this needs not to be the case. When strong substitution effects are likely to occur after a quality change for even only one of the goods, a multiple good framework is recommended.

3. Methodology: a dynamic empirical model

[1] The proof is given by Foster and Just (1989) and is based on $e(p_1,U_0,\boldsymbol{\theta}_1)$ having the same (x,z) solution of the constrained cost function. The unconstrained function then must be corrected to account to the fact that the price of x is actually p_0 rather than p_1.

The original work on COI by Foster and Just (1989) dealt with a milk contamination case occurred in the Hawaiian Isles, where information was upheld for a time. Subsequently, Teisl and Roe (1998) found COI a suitable measure in order to assess the benefit of a labelling policy. An empirical application by Teisl et al. (2001) assessed the welfare impact of nutritional information in the context of an experimental labelling program. All studies used a revealed preference approach based on estimation of demand curves, which allows to retrieve the cost function and the COI measure. Foster and Just retrieved the cost function from a single equation Marshallian demand function, following the Hausmann method (Hausmann, 1981). Conversely, Teisl et al. estimated a demand system theoretically derived within a cost minimisation framework. As mentioned in the previous section, the latter approach is less restrictive than the former as it allows for substitution effects among related goods.

However, retrieving a cost function and undertaking welfare measurement raises some methodological issues.

First of all, basing welfare measures on cost functions estimated from systems of demand equations is theoretically correct only for complete demand system taking into account all goods consumed by households. However, most estimated systems are not complete ones, due to data limitations. In empirical works, weak separability is often invoked to allow estimation of demand system for a set of goods conditional on the total expenditure on the group. Unfortunately, cost functions recovered from conditional (also known as partial) demand systems do not provide unbiased welfare measures[2].

Conversely, La France and Hanemann (1989) demonstrate that under fairly general condition welfare measure derived from incomplete demand system are unbiased. In order to define an incomplete demand system, let us consider a set of goods whose price vector **p** is known and a residual set of goods associated with the – generally unknown – price vector **q**.

If the system is integrable - as it would be the case under fairly general conditions - then it exists a cost function which is well behaved with respect to **p**, but not jointly to (**p**,**q**):

$$e(\mathbf{p},\mathbf{q},U) = \varepsilon\left[\mathbf{p},\mathbf{q},\eta(\mathbf{q},U)\right] \qquad (9)$$

where $\eta(\mathbf{q},U)$ is an arbitrary constant of integration that does not depend on **p** and whose structure is not recoverable.

Given that

$$\partial\eta(q,U)/\partial p_j = 0 \qquad (10)$$

the cost function recovered from the incomplete system reveals the welfare effects of changes in **p.** Hence, estimation of incomplete demand systems is an interesting option if one has data on x, **p**, **q** and θ and wishes to derive a CV measure for a change in the subset of the prices in **p** (Hanemann and Morey, 1992).

A second methodological issue is raised by the need to account for dynamics, due to the frequent empirical finding of serially correlated residuals from demand system models estimated on time series data[3]. In this work we rely on the flexible cost function provided in the context of the incomplete Almost Ideal Demand System (AIDS) (Deaton, Muellbauer, 1980). Several dynamic

[2] It has been demonstrated by La France (1993) and Hanemann and Morey (1992) that CV measures derived from conditional demand systems correspond to true CV only when consumption of the excluded goods is assumed to be fixed at the initial level. Furthermore conditional CV provides only a lower bound for the true welfare change.

[3] Indeed, Deaton and Muellbauer themselves remarked that serial correlation may arise due to the imposition of homogeneity constraints when the expenditure on certain items is inflexible in the short run, which is quite likely when working with high-frequency (monthly or weekly) data. The economic meaning of lagged expenditure shares entering the model has been explained by "myopic" habit persistence behaviour (as e.g. in Pashardes, 1986).

structures in the AIDS model framework have been proposed (see Ray, 1984, Alessie and Kaptein, 1991, Attfield, 1997 among others). Here we refer to a partial adjustment structure, as in Kesavan et al. (1993), Rickertsen (1996) and Edgerton (1996).
The equations of the incomplete partial adjustment AIDS model for *n-1* items and a numeraire good (such as the residual composite item) are specified as follows:

$$w_{it} = \alpha_{it}^{*} + \sum_{j=1}^{n-1} \gamma_{ij} \log p_{jt} + \beta_i \log\left(\frac{y_t}{P_t}\right) \qquad i: 1,\dots, n \qquad (11)$$

where w_{it} is the expenditure share for item i at time t; p_{jt} is the price of item j at time t, deflated by the price of the numeraire good p_{nt}; y_t is the total per capita expenditure at time t, also deflated by p_{nt}. The price of the numeraire good - usually a price index for the composite item - enters the system just as a deflator for prices and expenditure. This deals with the lack of data about the price vector of the goods excluded from the incomplete system (**q**): as in LaFrance (1993). Following Deaton and Muellbauer, the nonlinear price index P_t is defined as follows:

$$\log P_t = \alpha_0 + \sum_i \alpha_{it}^{*} \log p_{it} + \frac{1}{2}\sum_{i=1}^{n-1}\sum_{j=1}^{n-1} \gamma_{ij} \log p_{it} \log p_{jt} \qquad (12)$$

The intercept for the *i*-th equation, accounting for monthly seasonality, information and dynamics, is given by the following expression:

$$\alpha_{it}^{*} = \sum_{s=1}^{12} \delta_{is} a_{st} + \sum_{j=1}^{n-1} \psi_{ij} w_{j,t-1} + b_i d_{zt}^{*} (1 + \log t_z^{*})^{c} \qquad i: 1,\dots,n \qquad (13)$$

The first addendum on the right hand side of (13) represents the seasonal intercept, with a_{st}=1 when the *t*-th observation falls in month *s* and a_{st}=0 otherwise. The second addendum allows for dynamic behaviour by entering the vector of lagged expenditure shares.
Finally, the last addendum is a nonlinear intercept shift (similar to the one proposed by Foster and Just, 1989) which allows to embody the effects of retained information[4]. Such intervention structure implies that the sign and relevance of the disclosed information on the *i*-th intercept depends upon the item-specific coefficient b_i and evolves over time according to a parameter *c* which is held constant across equations. The *a priori* information about the timing of the disclosure of information (say time *z*) enters the system through d_{zt}^{*} which equals to 1 if *t*□ *z*. By setting t_z^{*} =*t-z-1* we allow the impact of disclosed information to evolve over time with a direction and a rate which depend upon *c*, where *c=0* implies a constant effect, *c<0* a decreasing effect and *c>0* an increasing effect. For *c<0* the intervention tends asymptotically to 0. The evolution over a ten years time span according to a set of values of *c* is shown in figure 3[5].

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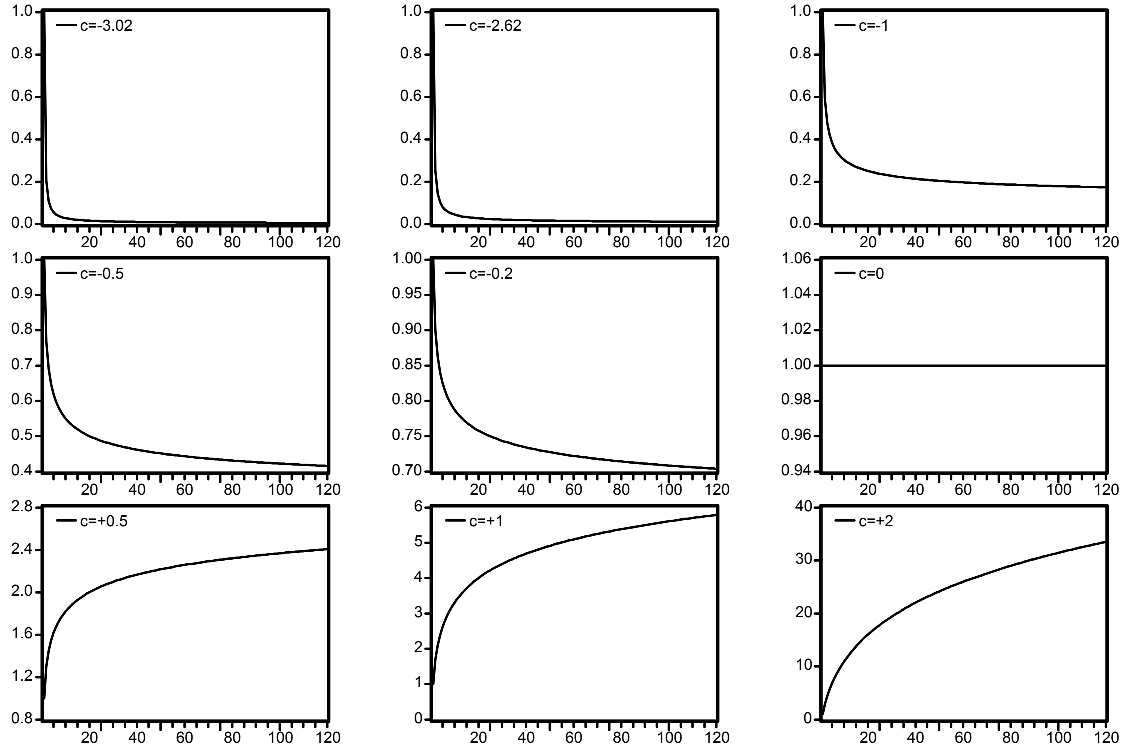
[4] Alternatively, if data are available, one can define the shift in demand as a function of “positive” and “negative” information released by the media (as in e.g. Liu et al., 1998 and Verbeke et al., 2000).
[5] It can be shown analytically that the shift becomes negligible (<0.01) in ten years when c<-2,62, in five years when c<-2.83 and in three years when c<-3,02

Figure 3. Value of the intervention variable over 120 months for different values of c.



The system in (11) is nonlinear due to nonlinearities in the price index P_t and in the modified intercept α_{it}^{*}. The problem of singularity is not present here, as the demand system is incomplete and the numeraire item equation is dropped by construction. As the lagged expenditure shares are pre-determined, these can be considered as exogenous and (11) is still a seemingly unrelated regression (SUR) system. We exploit a two-stages method for obtaining maximum likelihood estimates similar to that exploited by Deaton and Muellbauer (1980). In the first stage, the system coefficients are estimated by iterating the Zellner (1962) SUR estimator conditional to the value of the nonlinear price index P_t. In the second stage, the index P_t is explicitly computed using the coefficient estimates obtained in the first stage. This procedure is iterated until convergence, which is achieved when the estimates of the system coefficients are stable[6].

The COI measure can be derived from the coefficient estimates, according to (8). The first step consists in computing the vector of prices $\mathbf{p_{1t}}$ as in (7) for each time period following the crisis. As mentioned, the vector $\mathbf{p_{1t}}$ represents the price that would need to be charged to cause the individual to purchase at time *t* the same quantity he would have chosen before the release of information, keeping the utility level constant at the level U_0. This can be derived by holding fix the quantities[7] and the total expenditure at their value at the time $r=z-1$, i.e. the last time period before the release of information, and solving numerically the system (11) with respect to prices[8]. Once $\mathbf{p_{1t}}$ has been computed, it becomes possible to estimate the COI measure by exploiting the AIDS model cost function. For the system in (11), given that the utility level is U_0, after inversion of the utility function, the cost functions can be written as follows:

[6] The final estimates will be maximum likelihood estimates, as the Oberhofer-Kmenta conditions are met for the SUR model (Greene, 1997, p. 681).

[7] Defined as the ratio between expenditure and price for each good

[8] This involves a nontrivial dynamic optimisation problem, as lagged expenditure shares are considered.

$$\log e(\mathbf{p_1}, U_0, \boldsymbol{\theta_1}) = \alpha_0 + \sum_{i=1}^{n-1} \alpha_{it}^* \log p_{1it} + \frac{1}{2} \sum_{i=1}^{n-1} \sum_{j=1}^{n-1} \gamma_{ij} \log p_{1it} \log p_{1jt} + \\ + \prod_{j=1}^{n-1} \left(p_{1jt} / p_{jr} \right)^{\beta_j} \left\{ \log y_r - \alpha_0 - \sum_{i=1}^{n-1} \alpha_{it}^* \log p_{ir} - \frac{1}{2} \sum_{i=1}^{n-1} \sum_{j=1}^{n-1} \gamma_{ij} \log p_{ir} \log p_{jr} \right\} \quad (14)$$

$$\log e(\mathbf{p_0}, U_0, \boldsymbol{\theta_1}) = \alpha_0 + \sum_{i=1}^{n-1} \alpha_{it}^* \log p_{ir} + \frac{1}{2} \sum_{i=1}^{n-1} \sum_{j=1}^{n-1} \gamma_{ij} \log p_{ir} \log p_{1jr} + \\ + \left\{ \log y_r - \alpha_0 - \sum_{i=1}^{n-1} \alpha_{it}^* \log p_{ir} - \frac{1}{2} \sum_{i=1}^{n-1} \sum_{j=1}^{n-1} \gamma_{ij} \log p_{ir} \log p_{jr} \right\} \quad (15)$$

where p_{1it} is the i-th element of the vector $\mathbf{p_{1t}}$, computed as previously described and p_{ir} are the actual prices for the i-th good at time r, i.e. the last period prior to the information release. From (14) and (15) is it possible to compute a measure of the cost of ignorance:

$$COI_t = e(\mathbf{p_1}, U_0, \boldsymbol{\theta_1}) - e(\mathbf{p_0}, U_0, \boldsymbol{\theta_1}) + \sum_{i=1}^{n-1} \left(p_{iz} - p_{1it} \right) q_{iz} \quad (16)$$

where $q_{it} = \frac{w_{it}^* x_t}{p_{it}}$ are the quantity indices and w_{it}^* are the fitted values for the expenditure shares, computed through (11).

4. An application to the BSE scare in Italy

The empirical model detailed in the previous section was applied to the case of BSE crisis in Italy. The event of interest was the release of information about a possible linkage between BSE and the human disease CJD at the end of March 1996. Total per capita household expenditure and consumer expenditure on the following four grouped foods was considered: beef, poultry, other meats, fish. Aggregate (monthly) data were constructed from individual household data drawn from the ISTAT Italian Household Expenditure Survey over the period 1986-1999. Nominal price data were built by using the nominal 1996 price data from the ISMEA-Nielsen household budget survey and the time series of the ISTAT price indices. The price index for the residual (numeraire) good was the ISTAT Consumer price index.

Results of the AIDS Estimation and information impact

Model estimates are reported in table 1. As illustrated in the previous section, the non linear intercept shift which allows to embody the effects of information depends on both an item specific coefficient b_i and a common parameter c acting as exponent of a function of time elapsed from the release of information[9]. The item specific coefficient resulted to be significant – with the expected negative sign – for the beef equation, whereas the shifts on the other equations did not emerge as statistically significant. The c parameter, shaping the evolution of the effects, was -0.86,

[9] Conversely to Foster and Just, in (14) we suggest to use the logarithm of t, which involves a more gradual evolution of the impact.

corresponding to a decreasing impact of information. The estimated pattern for each equation is reported in figure 4.

Table 1. Model estimates and diagnostics

	Beef	Poultry	Other meats	Fish
Intercept (avg)	-0.12	**-0.13**	**0.22**	**0.22**
	(0.10)	*(0.03)*	*(0.09)*	*(0.08)*
$\psi 1$	**0.488**	**0.056**	0.112	0.021
	(0.07)	*(0.02)*	*(0.06)*	*(0.06)*
$\psi 2$	-0.313	-0.009	-0.081	-0.136
	(0.28)	*(0.09)*	*(0.26)*	*(0.22)*
$\psi 3$	0.170	0.011	**0.240**	0.026
	(0.09)	*(0.03)*	*(0.08)*	*(0.07)*
$\psi 4$	0.016	0.010	**0.388**	**0.556**
	(0.12)	*(0.04)*	*(0.10)*	*(0.09)*
b	**-0.0065**	0.0004	0.0001	-0.0004
	(0.0013)	*(0.0004)*	*(0.0012)*	*(0.0011)*
c	**-0.863**	**-0.863**	**-0.863**	**-0.863**
	(0.22)	*(0.22)*	*(0.22)*	*(0.22)*
$\gamma 1$	-0.0003	0.0006	-0.0004	-0.0113
	(0.012)	*(0.004)*	*(0.011)*	*(0.010)*
$\gamma 2$	0.0151	0.0029	**0.0192**	-0.0003
	(0.008)	*(0.002)*	*(0.007)*	*(0.006)*
$\gamma 3$	0.0062	0.0105	-0.0250	-0.0047
	(0.020)	*(0.006)*	*(0.018)*	*(0.016)*
$\gamma 4$	**0.0143**	**0.0049**	-0.0046	0.0042
	(0.005)	*(0.001)*	*(0.004)*	*(0.004)*
β	**-0.016**	**-0.003**	**-0.008**	**-0.008**
	(0.002)	*(0.001)*	*(0.002)*	*(0.002)*
R-square	0.96	0.91	0.80	0.52
D.W.	2.45	1.98	2.42	2.11

Note: Standard errors in brackets; value significant at a 0.01 confidence level are reported in bold.

Table 2. Price and expenditure elasticities before and after March 1996. Estimates of the Dynamic demand system from monthly data (1986-99).

	Beef	*Poultry*	*Other meats*	*Fish*	*Expenditure*
			Before March 1996		
Beef	-1.18	0.35	0.08	0.34	0.51
Poultry	0.09	-0.65	1.12	0.56	0.69
Other meats	-0.02	0.88	-2.06	-0.16	0.63
Fish	-0.56	0.16	-0.09	-0.55	0.53
			After March 1996		
Beef	-1.26	0.51	0.12	0.49	0.29
Poultry	0.11	-0.57	1.38	0.69	0.62
Other meats	-0.02	0.90	-2.09	-0.16	0.62
Fish	-0.54	0.16	-0.09	-0.57	0.54

The intervention on beef, poultry and other meats show the expected sign and trend, whereas fish shows a negative shift like beef. However, it is worth emphasising that the intervention is negligible for all equations but the beef one. These results show that the impact on beef consumption is negative and decreasing in absolute value as expected. Once trust in meat safety has been lost, it takes a long time to recover given the asymmetric impact of alarming and reassuring news. According to the estimated value of c, the impact of the crisis on beef demand is permanent, as the BSE impact becomes negligible only asymptotically[10].

Price and expenditure elasticity estimates computed from the appropriate coefficients are reported in table 2. Mean values for the period before and after the release of the BSE news are shown. Coherently with the above results, little change emerge in price elasticities for the poultry, other meats and fish equations. Instead, for what concerns beef demand, there is an increase in elasticity with respect to own price and cross prices. Expenditure elasticities are basically unchanged for other meats and fish, whereas there is a dramatic decrease in expenditure elasticity for beef and a less prominent one for the poultry equation. These results are consistent with other studies on the impact of the BSE crisis on meat demand (e.g. Burton and Young, 1996).

Cost of Ignorance Estimates

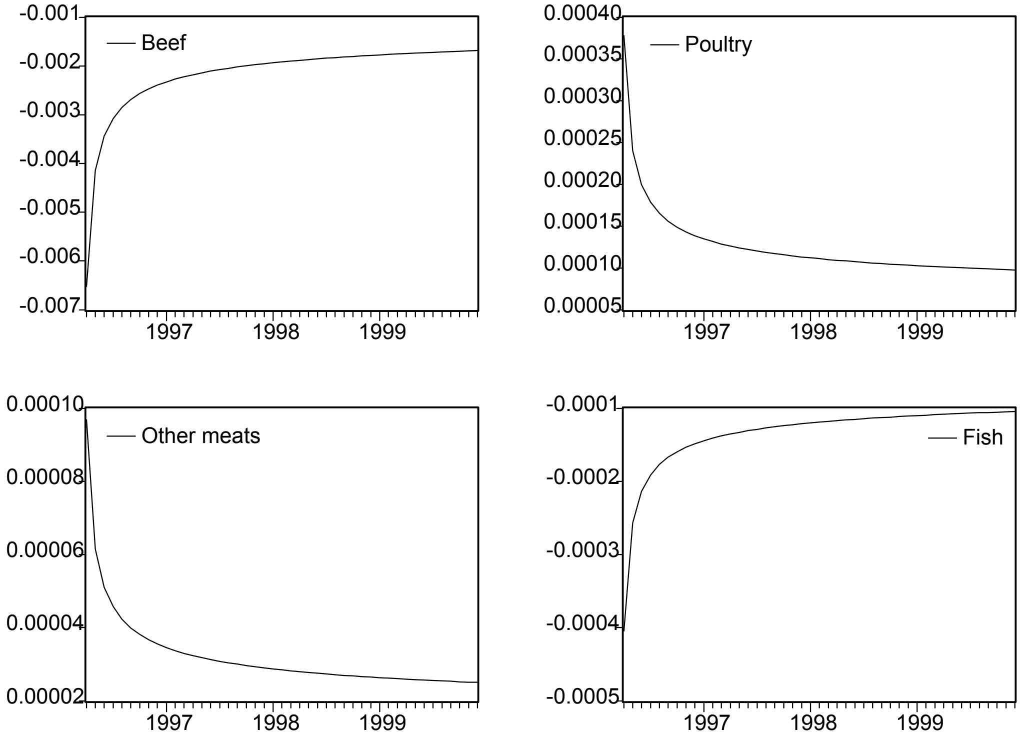
Following Foster and Just, the cost of ignorance for the four groups of meat and a numeraire is calculated according to (17). With regard to the BSE crisis, COI is a measure of the welfare improvement that could have been obtained if the news about the crisis had been released earlier. In particular, an earlier warning would have allowed consumer to reallocate their consumption across the foods of the meat group as they actually did once the BSE crisis was publicised. It is worth to note that, having included lagged expenditure shares in the cost function, the measure we obtain account also for the cost of adapting to new consumption patterns that differs from the habitual ones.

As a baseline for the COI measure we take the informational regime of February 1996, which is assumed to represent a state of ignorance. By that months the scare about BSE had not started yet and consumers believed consuming beef was safe as usual. However, some reports indicate that public authorities may have been aware of risk from BSE as early as 1988, when the first studies on transmissibility began.[11] The crisis began on 20 March 1996 and fully developed its effects on consumption since the following month. Taking February as a baseline, the COI for that month can be measured with respect to different informational status as reflected by the beliefs of consumers.

[10] This has little relevance, provided that a further relevant BSE crisis has affected Italian meat consumption in late 2000.

[11] In 1990, after evidence of transmission to cats, mice, pigs and sheep, public concern rose consistently. Despite the risks for human health could not be ruled out, all the governmental reactions were aimed to reassure the consumer about the safety of beef until March 1996 (source: BSE Inquiry, 2000. http://www.bseinquiry.gov.uk/).

Figure 4. Values of the estimated intervention variable for each food group (Apr 96-Dec 99)



The estimates of the intervention parameters suggest an immediate strong impact of information release on consumption, followed by a relatively slow decline. Indeed, as it was illustrated in section 2, subjective beliefs about the safety of beef change over time for a number of reasons. The original overreaction to bad news had probably influenced the sharp drop in consumption in the aftermath of the crisis. However, the slow recovering of consumption may have depended not only on a more rational reassessment of the actual risk, but also on the change in objective safety level following the adoption of countermeasures by the public authorities.

The choice of a true information reference month for COI calculation depends on which month after the crisis reflects subjective beliefs about safety level of beef that are as close as possible to the objective safety level of February 1996. Table 3 presents COI estimates for selected true information reference months. If, for example, we assume that beliefs in April 1996 are as close as possible to what was the objective safety level in February, a COI of € 22.05 per person per month is obtained[12]. This figure amounts to about 53% of the actual expenditure on the meat group foods.

Conversely, if a subsequent reference month - such as October 1996- is chosen, the estimated value of COI (€ 11.94) almost halves with respect to the previous figure. Thus the loss figure shows a decreasing trend as later reference months are chosen. However, the rate of decrease lowers as well so that the loss almost stabilises at a level of about € 7.5 by mid 1999. In all cases the estimate seems a plausible one. To draw a comparison, Foster and Just investigated a case of retained information about milk contamination in Ohau (Hawaii) for which they obtained a COI higher than the value of milk normally consumed. However, the present work differ from that of Foster and Just. Basing the COI estimates on a demand system rather than a single equation allows for substitution effects. In other words, conversely to the leftward shift for the Hicksian demand curve for beef observed in figure 2, the demand curve for beef substitutes perceived as safer after the release of information could shift rightward. If informed on the risk of beef, the consumer could have purchased the same quantities of substitute meats at higher prices. Hence, the COI measure in

[12] The measure is calculated at 1995 prices.

(9) also accounts for effects of information acting in the opposite direction and the system based measure is more appropriate than the single equation one.

Table 3. Measures of the per capita Cost of Ignorance (April 1996-December 1999)

Month	COI (€)	Percentage of per capita expenditure on meat and fish (Feb. 1996)
Apr-96	22.05	52.71
Oct-96	11.94	28.56
Apr-97	9.91	23.70
Oct-97	8.95	21.41
Apr-98	8.38	20.04
Oct-98	7.97	19.04
Apr-99	7.68	18.36
Oct-99	7.43	17.75

5. Conclusions

This paper has applied to the BSE scare the approach by Foster and Just for the measurement of consumer welfare losses associated with withheld information about food safety. A cost of ignorance (COI) is computed comparing the value of the cost function associated with the informed choice of food in the meat group with the value related to the uninformed one, assuming consumer are aware of the crisis. Unlike previous work, based on a single equation demand model, COI is measured retrieving a cost function from a dynamic Almost Ideal Demand System. This allows for reallocation of expenditure among foods of the same meat group, a behaviour actually observed in the aftermath of the BSE scare.
The COI measure computed here may prove useful to determine how much the Government should spend in providing timely information or to quantify the damage attributable to those who did not disclose the information in due time. The results of our analysis show that Italian consumers bore a sizeable loss because of the delayed release of information about BSE linkage with vCJD. The loss estimate per person per month ranges from 18% to about 50% of the total expenditure on the meat group depending on which period is assumed to embody the correct beliefs about the safety level of beef.
Further methodological refinements could improve the accuracy of the measure. Lack of data has prevented to model the subjective belief about the safety of beef as a function of the flow of positive and negative information that flooded consumers during the BSE scare. Alternative parameterisations of the intervention variable should be explored as well to check the robustness of results.
By providing an estimate of the welfare losses induced by mismanagement of information, our work allows a first insight into the theoretical and methodological intricacies of the assessment of the broader category of information policies. The BSE scare has shown the importance of risk communication in maintaining consumer's trust in the food chain and the European Union has recognised the importance of this aspect in its White paper on Food Safety. As the interest for information policies is growing, further studies are needed in order to provide sound economic basis for their assessment.

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