Health information and the demand for meat in Spain

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Summary

This paper analyses whether the growing amount of information about the relationship between diet and health has had an impact on the demand for different types of meat and fish in Spain. To achieve this objective, a health information index, based on the number of papers published in the MEDLINE database, is introduced into a 'CBS' system of demand equations. Given the time series properties of the variables, a cointegrated CBS model is estimated. Meat demand and health information elasticities are calculated. Results indicate that, in the case of Spain, health information elasticities are significant, having a positive effect on fish and poultry and a negative effect on beef and pork.

Keywords: meat consumption, health information, demand system, cointegration

JEL classification: Q11, C32, D12

1. Introduction

The relationship between diet and health has received increasing attention from health professionals, agricultural and food producers, consumers and policy-makers. Consumers' attention has been drawn to this relationship by the mass media and/or visits to doctors or nutritionists. In addition, the public sector has made considerable efforts to promote the advantages of healthier diets among consumers. It is not unrealistic to think that the increase in the availability of dietary health information has had an effect on the pattern of food demand.

In the last 20 years, various studies have evaluated the impact of health information on the demand for food (e.g. Brown and Schrader, 1990; Capps and Schmitz, 1991; Yen and Chern, 1992; Chern and Zuo, 1995). Most of these studies found that information on the health risks from a cholesterol-rich diet has induced significant changes in the consumption of certain food products. However, other studies, such as those by Chalfant and Alston (1988) and Robenstein and Thurman (1996), have shown no

evidence of such effects among meat products. For Europe, Rickertsen and von Cramon-Taubadel (2000) found mixed evidence of these effects in five European countries (France, Germany, Norway, Spain and the UK). Kim and Chern (1997) had already pointed out that contradictory results may arise from the problem of how health concerns are measured, that is, how the health index is constructed. Differences in sample periods and model specification may also be important causes of the contradictory results.

The objective of this paper is to assess the impact of health information on the demand for meat and fish in Spain. Two important methodological issues that have received attention in recent literature are featured in this paper. First, although the construction of any health information index is far from perfect, in this paper a realistic allowance for the carry-over effects of published information is used following Chern and Zuo (1995). Second, as time series data are used, the stochastic properties of the series included in the demand system are explored. The Johansen and Juselius multivariate cointegration framework is used and the health information index is assumed to be an exogenous variable.

The paper is organised as follows. Section 2 presents some descriptive data on intake levels of those foods more closely related to health information messages as well as the structure of meat and fish expenditure in Spain. In Section 3, theoretical and econometric frameworks are explained. After describing the data used, the main results are presented. Finally, the paper ends with some concluding remarks.

2. Demand for meat in Spain

Most of the medical literature on the relationship between health and diet deals with the importance of reducing fat and cholesterol intake so as to reduce the risk of food-related disorders such as heart disease, cancer, stroke and diabetes. In most developed countries, such information is highly relevant, as fat and cholesterol intakes surpass those recommended by health organisations such as the WHO (World Health Organisation). In Spain, a publication by the Ministry of Agriculture, Fisheries and Food (Ministerio de Agricultura, Pesca y Alimentacion, 1999) shows that both fat and cholesterol intakes greatly exceed the recommended levels. Also, the high level of fat intake is causing an important dietary imbalance in terms of different energy sources.[1]

The publication mentioned above indicates that, as far as the Spanish diet is concerned, the main sources of fat are meat, oils and dairy products, whereas cholesterol intake comes mainly through meat, eggs and fish. Taking into account the increasing information linking health to fats and cholesterol, on one hand, and to the main fat and cholesterol providers, on the other, it could be expected that the increase in health information would mainly affect meat consumption. However, as fish is considered a strong substitute

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1   A balanced diet should provide between 25 and 30 per cent of total calorie intake from fat whereas in Spain this percentage is around 46 per cent.

Table 1. Structure of Spanish meat expenditure (percentages of aggregate)

	1985	1989	1993	1996
Beef	36.3	37.1	33.2	31.2
Pork	15.6	14.5	15.0	16.0
Poultry	18.1	17.1	15.9	17.6
Fish	30.0	31.3	35.9	35.2

Source: own calculations from the Spanish Quarterly Household National Expenditure Survey.

for meat in Spain (Gracia and Albisu, 1995), our analysis should be carried out within a framework that jointly considers both meat and fish.

To form expectations about reactions of different classes of meat and fish demand to the increasing health information appearing in Spain, it is necessary to consider the fat and cholesterol composition of those product classes. Conversion factors published by the Ministerio de Agricultura, Pesca y Alimentación (1999) indicate that 100 g of beef, pork, poultry and fish provide 21.0, 23.0, 9.7 and 2.6 g of fat, and 0.70, 0.60, 0.81 and 0.54 g of cholesterol, respectively. Hence, information on a direct link between health and fat intake should induce an increase in poultry and fish consumption whereas the demand for beef and pork should decrease. Similarly, in the case of cholesterol, one would expect to see a replacement of meat products (mainly poultry and beef) by fish.

The recent evolution of beef, pork, poultry and fish expenditure shares is shown in Table 1. As can be observed, beef and fish expenditure shares have been following opposite trends. Whereas the beef expenditure share has decreased during the period analysed, the relative importance of fish within total expenditure on meat and fish has increased. In the case of pork, the budget share decreased slightly during the second half of the 1980s, and has increased since then. A similar evolution has taken place in the case of poultry, although here the turning point is located later. Our study aims to investigate to what extent these trends are due to the influence of health information on consumers' decisions, once allowance is made for the traditional economic factors of income and prices.

3. Theoretical and econometric frameworks

Most of the literature analysing the effect of any type of information on food demand is based on one of the two following approaches. The first one was suggested by Stigler and Becker (1977), who proposed using 'information' as an input variable in household production functions. In their formulation, information enters the (derived) demand function for market goods as a separate shift variable together with prices and income (e.g. Verma, 1980). An alternative approach was suggested by Theil (1980) and used by Duffy (1987) and Selvanathan (1989), among others. In this case, information

variables are considered 'taste shifters' that affect marginal utility and, as a consequence, 'information' variables enter the model as price deflators. Although the choice between these approaches is, to an extent, subjective, Brown and Lee (1993), analysing the effects of advertising on demand for several fruit juice products, showed that Theil's approach can easily be obtained by imposing certain restrictions on the absolute-price version of the Rotterdam model in which the information variable enters as an exogenous variable. Therefore, if under the household-production approach the Rotterdam model is chosen as the basic specification, it is possible to say that the household-production approach embeds Theil's approach as a special case. This was the main reason put forward by Capps and Schmitz (1991) and Kinnucan *et al.* (1997), among others, for choosing the Rotterdam model as the framework for introducing advertising and/or health information variables into their analysis.

On the basis of previous literature, the absolute-price version of the Rotterdam model is also chosen in this paper as the basic specification which, after including the health information variable, is expressed as

$$w_i \, \mathrm{d} \log q_i = \theta_i \, \mathrm{d} \log Q + \sum_{j=1}^{n} \pi_{ij} \, \mathrm{d} \log p_j + \vartheta_i \, \mathrm{d} \log \mathrm{HI} \qquad (1)$$

where q_i and p_i are the quantity and the price for good i, respectively; HI is the health information variable;

$$\theta_i = w_i \frac{\partial \log q_i}{\partial \log x} = p_i \frac{\partial q_i}{\partial x}$$

is the marginal propensity to consume; x is the total expenditure;

$$\mathrm{d} \log Q = \mathrm{d} \log x - \sum_{j=1}^{n} w_j \, \mathrm{d} \log p_j = \sum_{j=1}^{n} w_j \, \mathrm{d} \log q_j$$

is the Divisia volume index, and

$$\vartheta_i = w_i \left(\frac{\partial \log q_i}{\partial \log \mathrm{HI}} \right).$$

The system defined in (1) has two important limitations. First, it assumes that marginal budget shares are constant. However, there is no strong *a priori* basis for this conclusion; various workers conclude that this assumption is a severe handicap that may limit the validity of the model (e.g. Gao and Spreen, 1994; Lee *et al.*, 1994; Gao *et al.*, 1995). To escape this dilemma, one can rework the marginal budget shares in (1), θ_i, substituting those derived from the Working (1943) model, $\theta_i = w_i + \beta_i$. In this case, neither the budget share nor the associated marginal share are constant with respect to income. Introducing the new θ_i in (1) and rearranging terms, the following model is obtained:

$$w_i(\mathrm{d} \log q_i - \mathrm{d} \log Q) = \beta_i \, \mathrm{d} \log Q + \sum_{j=1}^{n} \pi_{ij} \, \mathrm{d} \log p_j + \vartheta_i \, \mathrm{d} \log \mathrm{HI}. \qquad (2)$$

This result (2) was proposed by Keller and van Driel (1985) of the Dutch Central Bureau of Statistics (CBS). For this reason, this model is known as the CBS model.

The second important limitation in (1), which is carried over into (2), is due to the fact that the model is specified in terms of infinitesimal changes (approximated by finite changes for estimation purposes) and, as a consequence, it supposes a first difference specification of the associated level version of the model. Considering that most of the variables involved in a level demand system are likely to be non-stationary, a differential system such as the Rotterdam or the CBS system yields more reliable results than models specified in levels (e.g. the Almost Ideal Demand System or the Linear Expenditure System). Nevertheless, recent developments in non-stationarity and cointegration analysis point out that a specification in first differences will be appropriate only if the corresponding level variables are integrated of order one, $I(1)$, and they are not cointegrated (Banerjee *et al.*, 1993). As a consequence, and following Gao and Lee (1995), a more correct model is obtained by starting with the level version of the model and then testing for non-stationarity and cointegration. The version of (2) expressed in levels is

$$w_i(\log q_i - \log Q) = \alpha_i + \beta_i \log Q + \sum_{j=1}^{n} \pi_{ij} \log p_j + \vartheta_i \log \mathrm{HI}. \qquad (3)$$

Demand theory restrictions can then be directly applied to the CBS parameters in (3). In particular, we have:

adding-up: $$\sum_{i=1}^{n} \beta_i = 0, \quad \sum_{i=1}^{n} \pi_{ij} = 0, \quad \sum_{i=1}^{n} \vartheta_i = 0 \qquad (4)$$

homogeneity: $$\sum_{j=1}^{n} \pi_{ij} = 0 \qquad (5)$$

symmetry: $$\pi_{ij} = \pi_{ji} \qquad (6)$$

negativity: $$\text{for any } \psi \neq 0, \quad \psi' \Omega \psi \leq 0, \qquad (7)$$

where $\Omega = \{\pi_{ij}\}$. Hence, Ω must be a negative semi-definite matrix.

Finally, elasticities for the *i*th product are calculated as follows:

income elasticity: $$\eta_i = \frac{\beta_i}{w_i} + 1$$

compensated price elasticity: $$\xi_{ij} = \frac{\pi_{ij}}{w_i}$$

uncompensated price elasticity: $$e_{ij} = \frac{\pi_{ij}}{w_i} - \eta_i w_j$$

health information elasticity: $$\kappa_i = \frac{\vartheta_i}{w_i}.$$

If the series in model (3) are non-stationary and cointegrated, then the system reveals a long-run optimal consumption pattern for, in our case,

meat products. The estimation of system (3) by ordinary least squares results in parameter estimates that are superconsistent but have non-standard distributions, making it impossible to test whether the theoretical restrictions hold, as tests based on standard asymptotic results will have the wrong size. Several workers have developed alternative techniques for the estimation and inference for long-run equilibrium relationships (e.g. Johansen, 1988; Phillips and Hansen, 1990; Park, 1992; Johansen and Juselius, 1994). However, the cointegration approach has only recently been applied to demand analysis (see, e.g. Denbaly and Vroomen, 1993; Balcombe and Davis, 1996; Attfield, 1997).

The estimation of cointegrated demand systems is routinely carried out using Engle and Granger's (1987) two-step procedure. This approach, however, has been subject to criticism (Ben Kaabia and Gil, 2001). An alternative framework for estimating a demand system has been proposed by Pesaran and Shin (1999), who estimated a Vector Autoregressive (VAR) model using Johansen's cointegration procedure. This approach has the advantage that it does not depend on an arbitrary division of the variables between endogenous and exogenous, and that the number of long-run relationships is not assumed *a priori*, but can be explicitly determined. Another important advantage of this approach is that an economically meaningful interpretation of these cointegration relationships can be achieved by imposing restrictions on the estimated parameters.

Thus, using recently developed statistical tools for analysing cointegrated data, this paper tries to extend to a CBS model the framework developed by researchers mentioned in the previous paragraph for the specification and interpretation of the long-run structure of a demand system.

4. Data

The quarterly data used in this analysis come from the Spanish Quarterly Household National Expenditure Survey carried out by the Instituto Nacional de Estadistica. This survey includes consumption and expenditure for beef, pork, poultry and fish, from which prices can be derived as unit values by dividing expenditure by quantities consumed. The sample period is first quarter 1985 to fourth quarter 1997.

4.1. The health information index

One of the most important issues when analysing how health information affects consumer decision-making is to define a variable that measures how well and when consumers are informed about the relationship between diet and health. It is usually assumed that this information is originally published in medical journals and then divulged to the general public by medical or mass media professionals. It is also assumed that the number of professional journal papers published is directly proportional to the amount of health information disseminated to consumers. Using these assumptions, health information indices have been generated by counting the papers published

in roughly 4,000 major medical journals and, as a consequence, this approach can be considered a good measure of the information potentially available to the public on the link between diet and health.[2] The seminal study using this approach was carried out by Brown and Schrader (1990), who investigated how health information affected US shell egg consumption. Later, other studies have analysed information effects on other products: for example, Capps and Schmitz (1991) and Kinnucan *et al*. (1997) on the demand for beef, pork, poultry and fish; Yen and Chern (1992) and Kim and Chern (1999) on the demand for fats and oils; and Chern and Zuo (1995) on fresh milk demand.

Health indices used in earlier literature differ mainly on two points. First, the number of published papers identified depends on the keywords used to perform the search. Keywords obviously depend on the aim of the study and the specific products being considered. Second, once the papers have been identified, the health index may be derived by simple counting or by more complex methods using weighting techniques. For instance, Brown and Schrader (1990) used as keywords 'cholesterol and (heart disease or arteriosclerosis)' and discarded as irrelevant papers those focusing on smoking, alcohol abuse, etc. In this case, they simply added up the number of published papers. Then, each paper that provided evidence of a link added one unit to the index whereas each paper that provided evidence in the opposite direction subtracted one unit. The resulting index closely resembled a trend.

Later, Chern and Zuo (1995) suggested a new procedure for constructing the index. They added 'fat(s)' to the key words used by Brown and Schrader (1990) and they did not manually discard 'irrelevant' papers. In relation to the index construction procedure, they assumed that all papers had both carry-over and decay effects. They specified a weighted function based on the number of relevant papers to account for these effects. To estimate the weights, they tested with cubic and third-degree distributed lag functions and with various lifespans for papers (12 and 24 months). By doing that, they tried to overcome the main criticism of Brown and Schrader's (1990) procedure relating to the cumulative nature of the index. It seems more reasonable to assume that the influence of a paper will decline over time and may even vanish if the message is not repeated. Furthermore, Chern and Zuo (1995) did not distinguish between supporting papers (unfavourable

2 Although MEDLINE contains only papers written in English, it has been judged to be the most relevant database to use for Spain. Most of the outstanding Spanish medical research is published in journals included in this database. Moreover, Spanish nutritionists use this database as the main source of information. Finally, as the number of medical papers increases, the number of articles in the mass media also increases more or less at the same rate. In fact, Chern (2000) showed that using MEDLINE or, as an alternative, the number of articles published in the *Washington Post*, the resulting health information indices were similar. As information in the mass media flows rapidly, it is not unrealistic to assume that news published in international newspapers dealing with scientific issues is also published in the Spanish mass media. In any case, we have explored the possibility of creating a similar index using a Spanish database but the number of journals included each month is very low, with many months having zero records.
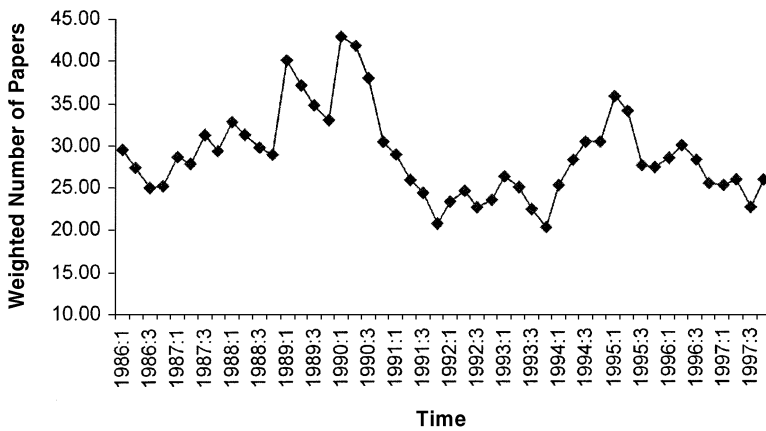
Figure 1. Health information index evolution.

information) and questioning papers (favourable information), as the number of questioning papers was rather small. This approach is also supported by workers such as Kinnucan and Chang (1993), who argued that consumers were especially responsive to 'negative' information and, thus, a simple subtraction rule may overemphasise 'positive' information.

In this paper, the health information index has been constructed following Slåen (1999), where the following key words are used: '(fat(s) or cholesterol) and (heart disease or arteriosclerosis) and diet'.[3] Then, Chern and Zuo's (1995) approach is used so as to include a third-degree distributed lag function and a lifespan of four quarters. The resulting health information index is shown in Fig. 1. As can be observed, unlike other indices, the pattern followed is different from a time trend.

5. Results

5.1. Non-stationarity

The order of integration of each variable in (3) was tested using both the Augmented Dickey–Fuller (Dickey and Fuller, 1981) and the KPSS (Kwiatkowski *et al.*, 1992) tests. Special attention was paid to the selection of lag length (truncation parameter in the KPSS test) and to the deterministic components included in the model. Results indicate that the hypothesis that all variables are $I(1)$ cannot be rejected at the 5 per cent level of significance.[4]

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3 Slåen (1999) found that although such keywords are not substantially different from those used by Chern and Zuo (1995), the number and relevance of papers found was higher. For instance, in 1994, Chern and Zuo (1995) found only 82 hits whereas Slåen's keywords resulted in 155 hits, most of them being relevant.

4 Results are not shown because of space limitations but are available from the authors upon request.

5.2. Cointegration rank

Taking into account that all the variables in (3) are $I(1)$, the Johansen and Juselius procedure is used to check the possible existence of stationary equilibrium relationships among them. The base-line econometric specification for multivariate cointegration is a VAR(p) representation of a k-dimensional time series vector Y_t reparameterised as a Vector Error Correction Model (VECM):

$$\Delta Y_t = \mu D_t + \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} - \Pi Y_{t-1} + e_t \qquad (8)$$

where: $Y_t = [\tilde{w}_{bt}, \tilde{w}_{pt}, \tilde{w}_{lt}, LP_{bt}, LP_{pt}, LP_{lt}, LP_{ft}, LQ_t, LHI_t]'$ is a $(k \times 1)$ column vector where $\tilde{w}_{it} = w_{it}(\log q_{it} - \log Q_t)$, with $i = b$, p, l (b, beef; p, pork; l, poultry). LP_{bt}, LP_{pt}, LP_{lt}, LP_{ft} represent prices in log form for beef, pork, poultry and fish, respectively. LQ_t is the logarithm of real expenditure and LHI_t is the logarithm of the health index. D_t is a vector of deterministic variables (intercepts, trend) where μ is the matrix of parameters associated with D_t. The Γ_i are $(k \times k)$ matrices of short-run parameters ($i = 1, \ldots, p - 1$), where p is the number of lags. Π is a $(k \times k)$ matrix of long-run parameters and e_t is the vector of independently and identically distributed disturbances characterised as $N(0, \Sigma)$.

In modelling the short-run and to reduce the number of estimated parameters, the health index (LHI) is treated as an exogenous $I(1)$ variable, which is an economically plausible assumption. In the terminology of Pesaran *et al.* (2000), the health index is considered a 'long-run forcing' variable in the determination of consumption behaviour. In other words, for the partition $Y_t = (Z_t, LHI_t)'$, the model (8) can be reparameterised assuming that changes in the health index have a direct influence on the variables $Z_t = [\tilde{w}_{bt}, \tilde{w}_{pt}, \tilde{w}_{lt}, LP_{bt}, LP_{pt}, LP_{lt}, LP_{ft}, LQ_t]'$ of the system, whereas LHI is not affected by the changes in the equilibrium relationships nor by past changes in Z_t. This is equivalent to the notion that the set of variables Z_t does not Granger-cause LHI_t.

Under these conditions the system for Z_t conditioned by the value of LHI_t can be written as

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \sum_{i=0}^{k-1} \Lambda_i \Delta LHI_{t-i} - \Pi Y_{t-1} + e_t. \qquad (9)$$

To test for the number of cointegrating vectors (r), we use the trace statistic (which tests whether there are at least r cointegrating vectors against the maintained hypothesis (rank $(\Pi) = p$)); and the λ_{max} statistic (which tests whether there are r cointegrating vectors against the alternative $r = r + 1$) (Johansen, 1988). If the hypothesis of cointegration is not rejected, Y_t is said to be cointegrated in the sense that there exists a $k \times r$ matrix λ such that $(\lambda' Y_{t-1})$ is stationary and, consequently, the cointegration relationships can be can formally expressed as $\Pi = \gamma \lambda'$. Each column of the matrix λ represents a cointegrating vector, whereas the rows of the γ matrix represent the adjustment coefficients that determine the speed of

Table 2. Univariate and multivariate misspecification tests for the estimated VAR(2) model

	$\tilde{w}_b$	$\tilde{w}_p$	$\tilde{w}_l$	LP_b	LP_p	LP_l	LP_f	LQ
Normality test[a]	2.55	0.18	2.56	3.78	0.19	1.81	1.46	0.71
ULM(1)[b]	0.01	4.76	1.16	0.09	1.19	3.02	2.00	5.43
ULM(4)[b]	0.21	4.90	3.26	2.75	3.56	0.20	1.73	0.91
Vector normality[c]	$\chi^2(16) = 8.78$		$P = 0.92$					
MLM(1)[d]	$\chi^2(64) = 61.50$		$P = 0.57$					
MLM(4)[d]	$\chi^2(64) = 78.60$		$P = 0.1$					

[a] Jarque–Bera normality test. The critical value at the 5 per cent level of significance is 3.84.
[b] Breusch–Godfrey univariate autocorrelation test of order $i = 1, 4$. The critical values at the 5 per cent level of significance are 3.84 and 9.49, respectively.
[c] Doornik (1995) multivariate normality test.
[d] Godfrey (1988) multivariate autocorrelation test.

adjustment of the $k - 1$ equations to equilibrium after a shock to a long-run relationship.

Before estimating the lag length in (9) and the cointegration rank of Y_t, we consider the deterministic components to be introduced in the model. Results from unit root tests in many cases indicated that the variables were non-stationary with non-zero means. This seems to indicate that the specification of a VECM with a restricted intercept in the cointegration space may be appropriate. The lag length of the VAR model is usually chosen according to the Akaike Information Criterion (AIC) or some other information criterion and by checking the properties of the estimated residuals. In our case, taking into account the reduced sample period, a VECM with two lags is considered an appropriate choice. In any case, the results from uni-variate and multivariate misspecification tests are reported in Table 2 and corroborate the appropriateness of this specification.

Having obtained a correctly specified model, we now determine the cointe-gration rank. This is usually tested by using the maximum eigenvalue (λ-max) and the trace test statistics proposed by Johansen (1988). As an exogenous $I(1)$ variable is introduced in the system, critical values are taken from Pesaran *et al.* (2000). Results are shown in Table 3. The λ-max statistic suggests two long-run relationships whereas the trace statistic indicates the presence of three cointegrating relationships. However, taking into account the relatively large dimension of the VECM and the small sample available, the outcome of the test procedure has to be interpreted with some caution. Several simulation studies (Johansen, 2000; Gredenhoff and Jacobson, 2001) show that the asymptotic critical values may not be very close approx-imations in small samples.

Taking into account such results, in this study we base the final decision on the roots of the companion matrix and on graphs of cointegrating vectors. The eigenvalues of the companion matrix show that the first five roots were

Table 3. Results from cointegration rank tests[a]

r	$p-r$	Trace	CV(95%)	CV(90%)	λ-max	CV(95%)	CV(90%)
0	8	251.89	182.73	176.42	67.65	55.26	52.23
1	7	184.24	147.98	142.17	62.3	49.57	46.59
2	6	121.94	116.3	110.5	41.05	43.76	40.93
3	5	80.89	86.58	82.17	34.69	37.48	34.99
4	4	46.2	62.75	59.07	16.45	31.48	29.01
5	3	29.75	42.4	39.12	15.68	25.54	22.98
6	2	14.07	25.23	22.76	7.86	18.88	16.74
7	1	6.21	12.45	10.5	6.21	12.45	10.5

[a]CV are critical values taken from Pesaran *et al.* (2000).

close to unity whereas the rest were small, confirming that the system appeared to be $I(1)$, probably with three cointegrating vectors and five unit roots. Moreover, Fig. 2 shows that the first three cointegrating vectors appear to be stationary. Therefore, the overall evidence supports the presence of three stationary equilibrium relationships. As we have considered four products, the number of cointegrating vectors is the same as the number of CBS equations to be estimated, as a result of the adding-up restriction. However, further checks are needed before these cointegrating vectors can be identified as long-run CBS equations.

5.3. Long-run identification of the CBS model

When there is more than one cointegration vector, the economic interpretation of the estimated coefficients is not straightforward: the estimated coefficient matrices α and β are not necessarily uniquely determined, as any linear combination of stationary vectors is also a stationary relationship. In those
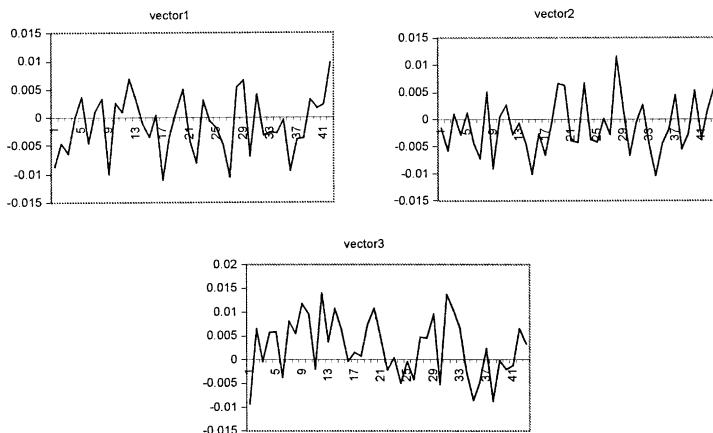
Figure 2. Cointegrating vectors.

cases, the economic interpretation of the cointegrating vectors as structural long-run relationships requires the imposition of at least r^2 restrictions (r of which are provided by normalisation conditions) on the cointegration space. These restrictions can be motivated by economic arguments (e.g. a particular variable does not appear in a specific vector, or homogeneity or/and symmetry restrictions). Johansen and Juselius (1994) have developed a testing procedure to identify the cointegrating vectors by imposing linear restrictions on them. Doornik (1995) and Pesaran and Shin (1999) generalised this procedure, allowing for the imposition of non-linear restrictions.

Let us start by considering the unrestricted cointegration space, given by

$$
\lambda' Y_t = \begin{pmatrix} \lambda_{11} & \lambda_{12} & \lambda_{13} & \lambda_{14} & \lambda_{15} & \lambda_{16} & \lambda_{17} & \lambda_{18} & \lambda_{19} & \lambda_{110} \\ \lambda_{21} & \lambda_{22} & \lambda_{23} & \lambda_{24} & \lambda_{25} & \lambda_{26} & \lambda_{27} & \lambda_{28} & \lambda_{29} & \lambda_{210} \\ \lambda_{31} & \lambda_{32} & \lambda_{33} & \lambda_{34} & \lambda_{35} & \lambda_{36} & \lambda_{37} & \lambda_{38} & \lambda_{39} & \lambda_{310} \end{pmatrix} \begin{pmatrix} \tilde{w}_{bt} \\ \tilde{w}_{pt} \\ \tilde{w}_{lt} \\ LP_{bt} \\ LP_{pt} \\ LP_{lt} \\ LP_{ft} \\ LQ_t \\ LHI_t \\ 1 \end{pmatrix}.
$$

$$(10)$$

As can be observed, long-run relationships are different from CBS equations in (3). Because we have three cointegrating vectors, the structural estimation of the long-run relationships requires the imposition of at least nine just-identifying restrictions, three of which are provided by the normalisation condition. The following just-identifying restrictions are imposed on (10):

$$\lambda_{11} = \lambda_{22} = \lambda_{33} = 1; \quad \lambda_{12} = \lambda_{13} = \lambda_{21} = \lambda_{23} = \lambda_{31} = \lambda_{32} = 0.$$

As a result, the three cointegrating relationships can be interpreted as long-run CBS equations. Finally, demand theory conditions (homogeneity and symmetry) can be directly tested as over-identifying restrictions using the procedures mentioned above. Taking into account (10), the homogeneity hypothesis can be formulated as follows:

$$\sum_{i=4}^{i=7} \lambda_{1i} = 0; \quad \sum_{i=4}^{i=7} \lambda_{2i} = 0; \quad \sum_{i=4}^{i=7} \lambda_{3i} = 0.$$

The likelihood ratio (LR) statistic for testing the three over-identifying restrictions is 7.09, which is below the 5 per cent critical value of $\chi^2(3)$ ($= 7.81$). Thus, the homogeneity hypothesis cannot be rejected. Furthermore, and to be consistent with economic theory, we have carried out a joint test of the symmetry and homogeneity restrictions. The symmetry hypothesis

Table 4. Long-run estimated parameters of the CBS model with homogeneity imposed[a,b]

$\tilde{w}_b$	$\tilde{w}_p$	$\tilde{w}_l$	LP_b	LP_p	LP_l	LP_f	LQ	LHI	Constant
1	0	0	−0.036	−0.200	0.227	0.009	−0.073	0.017	0.399
			(−1.80)	(−10.53)	(8.73)	(0.643)	(−3.17)	(3.40)	(6.44)
0	1	0	−0.06	0.103	−0.098	0.055	0.06	0.015	−0.157
			(−3.16)	(5.72)	(−3.92)	(3.929)	(2.73)	(3.00)	(−2.66)
0	0	1	0.056	−0.134	0.131	−0.054	0.12	−0.019	−0.197
			(2.44)	(−6.09)	(4.23)	(−3.38)	(4.44)	(−3.17)	(−2.78)

[a] t ratios are shown in parentheses.
[b] All variables are defined in equation (8).

requires the introduction of the following cross-equation restrictions:

$$\lambda_{24} = \lambda_{15}; \quad \lambda_{34} = \lambda_{16}; \quad \lambda_{26} = \lambda_{35}.$$

The LR statistic for jointly testing the homogeneity and symmetry restrictions is 37.95 (critical value at the 5 per cent significance level of $\chi^2(6) = 12.6$), which indicates that the null hypothesis is strongly rejected by the data.[5] The negativity condition has been analysed by calculating the eigenvalues of matrix Ω in (7). As the rank of this matrix is three $(n-1)$, the negative semi-definite condition is satisfied if one eigenvalue is zero whereas the rest are negative. In this case, results indicate that this condition is satisfied, and the calculated eigenvalues are: 0, −0.108, −0.437 and −0.851. The estimated coefficients of the three cointegrating vectors under the homogeneity restrictions are given in Table 4. From these parameters, it is possible to calculate the corresponding long-run elasticities. The estimated conditional elasticities (weak separability has been assumed) are shown in Table 5.

As expected (taking into account average budget shares), beef and fish are considered luxury products in relation to total meat and fish expenditure. Poultry and pork can be defined as necessities. The effect of health information on meat consumption is significant for all products except for fish. This result is different from that of Rickertsen and von Cramon-Taubadel (2000), who found no effect. The dissimilar procedure used here to calculate the

5 Several simulation studies have shown that the size distortion of the asymptotic LR tests for hypothesis testing on the cointegration vectors can be considerable in small samples, as it is in our case (Gredenhoff and Jacobson, 2001). Gredenhoff and Jacobson (2001) suggested the use of bootstrapping techniques to reduce size distortions. This method has proved more efficient than the traditional adjustment method used, for instance, by Reimers (1992). Using the latter approach as an example, the symmetry restriction is also rejected. However, the main problem we have found is that it has been almost impossible to estimate the system with the symmetry restriction imposed. This is one of the main limitations of the multivariate cointegration framework used in this paper, as the algorithms in existing software packages are insufficiently robust to obtain any kind of convergence once the over-identifying restrictions have been imposed. We have tried giving alternative initial values to estimated parameters and only in one case (that presented in the paper) have we reached convergence.

Table 5. Conditional long-run demand elasticities for meat in Spain[a]

	Budget share	Expenditure	Health	Price[b]			
				Beef	Pork	Poultry	Fish
Beef	0.346	1.211	−0.049	**−0.315**	0.578	−0.656	−0.026
		(18.35)	(−3.27)	(−9.05)	(10.51)	(−8.75)	(−0.65)
Pork	0.152	0.605	−0.099	0.395	**−0.770**	0.645	−0.362
		(4.17)	(−3.01)	(3.16)	(−8.02)	(3.93)	(−3.94)
Poultry	0.170	0.294	0.112	−0.368	0.882	**−0.821**	0.355
		(1.85)	(3.11)	(−2.44)	(6.08)	(−5.30)	(3.38)
Fish	0.332	1.322	0.039	−0.120	0.102	0.783	**−0.409**
		(6.09)	(0.81)	(−0.64)	(0.50)	(3.17)	(−6.71)

[a] t ratios are shown in parentheses.
[b] Marshallian own-price and Hicksian cross-price elasticities.

health information index as well as our explicit consideration of dynamics through cointegration analysis might explain such differences. Health information has a negative effect on the demand for beef and pork whereas it is positive for poultry and fish. However, although significant, the magnitude is relatively small.

All own-price elasticities are negative, being more elastic in the cases of poultry and pork. A possible explanation of this result is that for those two products quality is fairly homogeneous, so price changes may stimulate the consumption of other meats or fish. In the case of beef and fish there is more variability in terms of quality and prices. Beef price increases can provoke a higher demand for lower-quality beef with total beef consumption remaining constant.

Compensated cross-price elasticities indicate a certain degree of substitution amongst healthier products (positive health elasticity), on one hand, and amongst less healthy products (negative health elasticity), on the other. Table 5 shows that the cross-price elasticities fish–poultry and pork–beef are both positive. On the other hand, and in general terms, the cross-price elasticity is negative (complementary products) for products belonging to different groups (i.e. the pairs poultry–beef or pork–fish). Finally, a significant substitution effect is found between the less expensive meats (pork and poultry), which is consistent with previous expectations.

5.4. Short-run analysis

Now that the long-run structural model has been estimated, we will focus on the short-run analysis. As pointed out by Pesaran and Shin (1996) it is important that the analysis of cointegration is accompanied by some estimates of the speed with which the markets return to their equilibrium level, once shocked.

Table 6. Estimated adjustment parameters (γ)

	$\hat{\gamma}_1$	$\hat{\gamma}_2$	$\hat{\gamma}_3$	t values for γs		
$\Delta\tilde{w}_{bt}$	−0.36	0.23	0.21	−1.961	0.901	1.681
$\Delta\tilde{w}_{pt}$	0.35	−0.15	−0.30	4.352	−1.638	−2.927
$\Delta\tilde{w}_{lt}$	0.21	−0.31	−0.16	2.809	−1.25	−1.723
$\Delta\tilde{w}_{ft}$	−0.21	0.23	0.25	—	—	—

[a] All variables are defined in equation (8).

Traditionally, the speed of convergence towards equilibrium is represented by the weights (given by the estimated coefficients of the α matrix) that the cointegrating vectors have in each equation of the system. Table 6 reports the adjustment coefficients associated with the first three equations. As can be observed, beef (poultry) consumption seems to react only to a disequilibrium in poultry (beef) consumption, whereas pork consumption adjusts significantly to a disequilibrium in beef and poultry consumption. In addition, the degree of complementarity and substitution between the different goods can be analysed by looking at the sign of the adjustment coefficients α_{ij}. As α_{13} and α_{31} are positive (Table 6) and have the same magnitude, poultry (beef) consumption increases when beef (poultry) consumption is above its long-run equilibrium level. These results suggest that beef and poultry are complements. On the other hand, α_{21} and α_{23} are both negative and significant. This indicates that pork consumption is held down when beef and poultry consumption are above their equilibrium level, suggesting that pork is a substitute for the other two meats.

Finally, when analysing the short run, it is interesting to calculate the impulse response functions, that is, the response of each variable in the system to a shock in any of the other variables. As the aim of the paper is to determine the impact of health information on meat demand, we are going to focus on the impulse responses to a shock in the health information index. As traditionally demand analyses have focused on elasticities, in this paper we analyse the short-run dynamics of demand elasticities and how they converge to the long-run equilibrium.

One of the main problems in calculating impulse response functions is that the existing contemporaneous correlation among innovations makes it difficult to isolate shocks from a single variable. In this paper we have calculated the generalised impulse response functions (Pesaran and Shin, 1998); this approach avoids the problem mentioned above. Figure 3 shows the short-run behaviour of the elasticities of demand for the various meat products to a shock in the health information index.

Short-run health information elasticities converge to their long-run values, which are shown in Table 5. In the first two quarters the response is small. The higher impact takes place in the third quarter, decaying afterwards to the long-run equilibrium. We conjecture that the reason for this lagged response
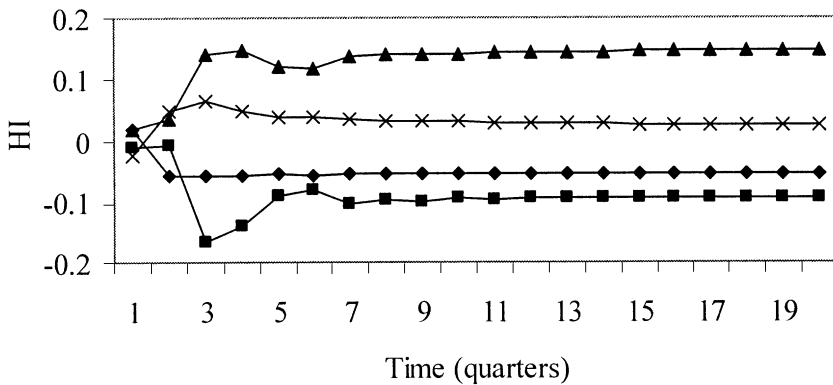
Figure 3. Short-run behaviour of the health information elasticity (HI) for different meat products. ◆, beef; ■, pork; ▲, poultry; ×, fish.

is twofold. First, and for us the most important reason, once the information is published it takes some time to reach consumers and, then, for consumers to adjust their behaviour. Second, taking into account the nature of the database, it is possible that information reaches Spanish consumers slightly later, although, as we have mentioned above when describing the MEDLINE database, we do not think this issue is relevant.

In the very short run, the negative impact is slightly higher in the case of beef whereas after three lags the negative effect on pork consumption is higher. The over-reaction for pork has to be analysed in conjunction with the poultry response, as they are close substitutes.

6. Concluding remarks

The increased information available to consumers about the relationship between diet and health affects consumer decision-making. Although most of this information is published in medical journals, the mass media and medical professionals are assumed to disseminate this information to consumers. As a consequence, consumers are aware that an appropriate diet reduces the risk of food-related diseases such as heart disease, diabetes and some types of cancer. The aim of this paper has been to analyse whether the increased information about the relationship between diet and health has had an impact on the demand for meats in Spain.

In pursuing this objective, three methodological issues have been discussed. First, recent developments have been followed in constructing an appropriate health index. Although no procedure is perfect, the approach followed here allows for carry-over and decay effects, which is more realistic than simply counting the number of papers supporting or questioning the health-and-diet relationship under analysis. As the resulting index does not behave like a time trend, it would be possible to distinguish health information effects from changing consumer tastes, an issue that could be explored in further research.

As time-series data have been used, a multivariate cointegration framework has been used to specify and estimate the cointegrated CBS model. The issue of estimating cointegrated demand systems has generated some discussion among applied economists. The procedure used in this paper is relatively new and, to our knowledge, this is the first time it has been applied to a model other than the Almost Ideal Demand System. However, the use of this procedure is not straightforward and has some limitations, mainly that in systems with more than four or five equations it is difficult to reach convergence when over-identifying restrictions are imposed.

Results obtained in this paper show a significant impact of health information on the demand for various meats. Although the magnitude of the impact is relatively small, it seems that increasing health information has a negative impact on red meats (beef and pork) and a positive impact on poultry and fish. It is therefore likely that the observed trends in Spanish fish and beef expenditures could be partially explained by the increasing concern about the relationship between diet and health. In the other two cases, the evolution of food consumption patterns is less clear.

At the same time, the reaction to available information is not instantaneous. Time is needed for the information to reach consumers, and consumers also take some time to adjust their behaviour. Further extension of this work is needed to cover a wider range of food products, to test alternative specifications for the health index and to provide alternative theoretical models for incorporating this information into demand analyses.

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