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Price transmission, BSE and structural breaks in the UK meat sector

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Abstract

This paper examines transmission between producer and retail prices for beef, lamb and pork in the UK and the impact of public concern over bovine spongiform encephalopathy (BSE) in early 1996. We used the cointegration procedure of Johansen *et al*. 2000), which admits structural breaks in cointegrating space. Results with monthly data for 1986–2000 show that a long-run relationship exists between each producer and retail price, and that a structural break occurs in the beef relationship at the height of the BSE crisis, which increases the margin by £1.12/kg. In contrast, there is no evidence of BSE-related breaks in the lamb or pork relationships.

Keywords: price transmission, BSE, UK meats, cointegration

JEL classification: Q13, D12

1. Introduction

The bovine spongiform encephalopathy (BSE) crisis in the UK in early 1996 had a profound effect on the meat sector. In the first two quarters of 1996, Meat and Livestock Commission (MLC) statistics show that the retail beef price fell by 11 per cent and consumption fell by 24 per cent; the retail lamb price increased by 23 per cent with consumption remaining largely unchanged; and the retail pork price increased by 2 per cent with consumption increasing by 4 per cent. Respective retail–producer price margins increased by 10, 59 and 12 per cent, and there was a public perception of profiteering. Our focus is on the long-run effects of the BSE crisis on producer–retail price transmission for beef, lamb and pork.

The literature on the economic impact of food scares is small. Among US studies, Brown and Schrader (1990) measured the effect of cholesterol information on the demand for eggs, and Kinnucan *et al*. (1997) examined the impact of health information on the demand for beef, pork, poultry and fish. In the UK, Burton and Young (1996) studied the effect of BSE on beef demand. The impact of food scares on price adjustments between different levels in the marketing chain has also received little attention. Lloyd *et al*. (2000, 2001a, 2001b) examined the impact of BSE on transmission between producer, wholesale and retail prices of beef using a three-equation vector autoregressive (VAR) model; evidence of cointegration in their beef model is weak until a fourth variable was added, namely a meat-scare index, which measures public awareness of BSE according to the number of newspaper articles in which it is mentioned. Commenting on the strengthened evidence of cointegration, Lloyd *et al*. (2001b: 356) concluded that it is consistent with pair-wise price co-movement and 'food publicity . . . plays a key role in the long-run evolution of UK beef price'. For lamb and pork, cointegration could only be found by including the meat-scare index and the beef marketing margin.

Our paper extends Lloyd *et al.* (2000, 2001a, 2001b) and provides further evidence on the existence of a long-run relationship between producer and retail prices for beef, lamb and pork using cointegration analysis.[1] The novel feature is our treatment of the effects on prices of changing consumer perceptions of risk from BSE. We allowed for the possibility that prices are subject to a shock caused by the BSE crisis, which in turn caused a structural break in the long-run relationship between producer and retail prices.[2] We used the recent cointegration procedure of Johansen *et al*. (2000), which permits up to two breaks in the cointegrating space at predetermined point(s) in time. Furthermore, we did not rule out that other exogenous events, such as the removal of the pound from the European exchange rate mechanism (ERM) on 'Black Wednesday' (16 September 1992), may also have caused structural breaks in price evolution. The paper is arranged as follows: Section 2 outlines a theory of marketing margins, Section 3 discusses our empirical methodology, Section 4 discusses the data and examines the results, and Section 5 concludes.

2. Some theoretical issues

The theory of marketing margins assumes simultaneous equilibrium at two market levels. The interaction of supply and demand at the retail level determines retail price; demand at the producer (farm-gate) level is derived from that at the retail level, which, with primary supply, determines producer price. The marketing margin (M) is the difference between the retail price (RP) and producer price (PP) per equivalent unit at equilibrium, and represents the price of marketing services such as processing, storage, wholesaling and retailing.

1 Our analysis, unlike that of Lloyd *et al*., did not include wholesale prices for three reasons: first, data on wholesale prices are less reliable than for producer or retail prices; second, there are alternative wholesale price series and none corresponds to either producer or retail prices in terms of reference quality; third, the wholesale prices for beef and lamb are sometimes less than producer prices, particularly before 1990.

2 Our analysis also extends the work of Tiffin and Dawson (2000), who use single-equation cointegration methods and test for a structural break in the UK lamb price relationship.

In Gardner's (1975) much-cited theoretical framework, a perfectly competitive marketing industry uses two inputs—an agricultural commodity and marketing services—to produce food. Retail food demand is determined by the retail price and an exogenous demand shifter (such as changing consumer preferences arising from concerns over BSE). Assuming profit maximisation and given the usual demand and supply conditions, equilibrium is unique. Gardner demonstrated the effect of an exogenous change in demand on the price spread measured by RP/PP: when the own-price elasticity of demand for a particular food is negative and the supply elasticity of the agricultural commodity is less than that of marketing services, as Gardner believed is likely, then the spread increases if the exogenous demand shifter falls.[3] Accordingly, we expected that the beef price spread would increase as a result of a change in preferences away from beef following the BSE crisis in early 1996. Gardner also showed that increasing marketing costs lead to a rise in RP/PP. During the BSE crisis, new meat regulations were imposed on the UK processing sector requiring new marketing inputs, which increased cost.[4] RP/PP therefore was expected to rise, reinforcing the initial increase caused by the shock to retail demand.[5] Furthermore, Gardner demonstrated that a shock to farm supply that reduces producer price increases RP/PP. Imperfect competition in the downstream food sector reinforces the widening of the price spread following a shock (McCorriston *et al.*, 1998, 2001).

George and King (1971) modelled the margin as a combination of an absolute spread (μ) and constant percentage of producer price (k). Under the hypothesis that margin behaviour depends on pricing practices in the upstream sector, $\mathrm{M} = \mu + k\mathrm{PP}$ and $\mathrm{RP} = \mu + \beta\mathrm{PP}$ where $\beta = 1 + k$. Gardner noted, however, that no simple mark-up pricing rule can accurately model the relationship between RP and PP. For instance, the variation in RP/PP at equilibrium following a shift in demand implies that a fixed percentage mark-up is not strictly appropriate. Thus, when estimating a bivariate regression between RP and PP, mixed effects result depending on which shocks (demand, farm supply or marketing services) dominate the time span.[6] We allowed for different constants when a shock occurs in the system so that $\mathrm{RP} = \mu_1 + \mu_2 + \beta\mathrm{PP}$ where μ_1 is the constant prior to a shock and μ_2 thereafter.

3 Gardner's (1975) model has been extended *inter alia* by Heien (1980), who incorporates sector dynamics, and Holloway (1991), who admits imperfect competition.

4 For example, the Specified Bovine Offal Order of 15 December 1995 prohibited the use of bovine vertebral column in the manufacture of mechanically reconstituted meat for human consumption.

5 The increase in RP/PP induced by the demand shift is due to a fall in PP in greater proportion than RP, whereas that induced by increasing costs leads *ceteris paribus* to an increase in RP with PP unchanged.

6 To address this limitation, Wohlgenant and Mullen (1987) introduced an interaction term between RP and the quantity moving through the marketing chain that allows prices to change as output changes.

3. Empirical issues

The first step was to establish the order of integration of the individual price series. The augmented Dickey–Fuller (ADF) test (Dickey and Fuller, 1981) is commonly used to test for unit roots but when a series is subject to a deterministic trend and an exogenous shock causes a structural break, the ADF test tends to under-reject (Perron, 1989). Accordingly, we used Perron's (1997) test where the null hypothesis is a unit root with a structural break and the alternative is stationarity around a broken trend (or level); the break point is estimated endogenously. Two types of break were examined: the first allows for a change in the level (intercept), and the second allows for changes in both level and trend (slope). Following Perron (1997), the test was performed by choosing the break point that minimises the t-statistic in the null, and the number of lags was chosen following the general-to-specific method of Said and Dickey (1984) where sequential F-tests were used at the 10 per cent significance level.[7,8]

If the series are I(1), the test for cointegration between RP_t and PP_t is a test of long-run equilibrium. If structural breaks within the two individual series occur, either at different times or at the same time and do not cancel each other out, cointegration must be analysed in a framework that allows for breaks in the deterministic components. Gregory and Hansen (1996) proposed a test of the null hypothesis (no cointegration) against the alternative (cointegration) allowing for a level shift, or a shift in both the level and the slope. However, this is not a formal test of the existence of a regime shift. Moreover, results depend on the normalisation chosen, and seasonal and short-run dynamics are neglected.[9]

Johansen *et al*. (2000) overcame these shortcomings and generalised the multivariate likelihood procedure of Johansen (1988) by admitting up to two breaks; two models were considered, one with a broken level and the other with a broken trend. We denote by q the number of periods into which the sample is divided, and each period by j. In the case of two breaks, the sample is divided into three periods ($j = 1, \ldots, q$ and $q = 3$) and the vector error correction model (VECM) is

$$\Delta Y_t = \alpha \begin{pmatrix} \beta \\ \mu \end{pmatrix}' \begin{pmatrix} Y_{t-1} \\ tE_t \end{pmatrix} + \gamma E_t + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + u_t, \qquad t = 1, \ldots, T \quad (1)$$

7 Perron tests were carried out in *RATS*, Version 5.02 (Estima, 2000) using the procedure, Perron97.src (http://www.estima.com/procs_unit.shtml#perron97).

8 With monthly data, both conventional and seasonal unit roots may be present. The latter are less plausible in economic time series, as they imply an evolving seasonal pattern. Hatanaka (1996: 15) expresses a 'doubt that seasonal variations are properly modelled by the (complex) unit roots. Deviations of seasonal fluctuations from the deterministic periodicity are bounded in probability at any time, and the (complex) unit roots are inappropriate to model the deviations.' For these reasons, we did not examine seasonal unit roots.

9 Hansen (1992) developed a test of the stability of cointegration vectors in a fully modified procedure but this does not identify which particular parameter is unstable. To our knowledge, this test has not been extended to the Johansen (1988) procedure.

where Δ is the difference operator; the vector $Y_t = (RP_t PP_t)$; p is the number of lags; E_t is a vector of q dummy variables $E_t = (E_{1t}E_{2t}\ldots E_{qt})'$ with $E_{j,t} = 1$ $(j = 1, \ldots, q)$ if observation t belongs to the jth period and zero otherwise, with the first p observations set to zero; and $D_{j,t-i}$ $(j = 2, \ldots, q$ and $i = 1, \ldots, p)$ is an impulse dummy that equals unity if observation t is the ith observation of the jth period. These dummies were included to render the corresponding residuals zero allowing the conditional likelihood function to be derived given the initial values in each period. The short-run parameters are Γ_i of order (2×2) for $i = 1, \ldots, p-1$, and $\kappa_{j,i}$ of order (2×1) for $j = 1, \ldots, q$ and $i = 1, \ldots, p$. The innovations, u_t, were assumed to be independently and identically normally distributed with mean zero, and symmetric and positive definite variance–covariance matrix Ω. The long-run drift parameters are $\mu = (\mu_1\mu_2\ldots\mu_q)$, α is a matrix of adjustment parameters, and β are the long-run coefficients in the cointegration vector. The cointegration hypothesis is formulated in terms of the rank (r) of $\pi = \alpha\begin{pmatrix}\beta\\ \mu\end{pmatrix}'$. The asymptotic distribution of the rank test depends on the number of non-stationary relationships, the location of the break points and the trend specification, and was determined by simulation with critical values from Nielsen (1999) and Johansen *et al.* (2000).

Johansen *et al.* (2000) examined two models corresponding to (1). In the first, there are no linear trends in the levels of the endogenous I(1) variables and the first-differenced series have a zero mean; here the broken level is restricted to the cointegration space. The second model does not account for long-run linear growth, and a broken linear trend is present in the cointegration vectors. To test between these models, the Pantula principle (Harris, 1995: 97) can be used to test the joint hypothesis of both rank and the deterministic components (Johansen, 1992).[10]

Once the cointegration rank and model are known, restrictions on the cointegration space can be tested using log-likelihood ratios (LR) as in the standard cointegration procedure. To illustrate, assume one cointegrating vector $(r = 1)$ with a broken level, and two breaks $(q = 3)$. Here, $(Y_{t-1}E_t)' = (RP_{t-1}PP_{t-1}E_{1t}E_{2t}E_{3t})'$ and the parameters in the cointegration vector are $(\beta\mu)' = (\beta_{RP}\beta_{PP}\mu_1\mu_2\mu_3)'$. Two hypotheses are of particular interest. The first is the null that each price belongs to the cointegration space. For example, the null of the exclusion of RP_t is

$$H_0^1\colon \beta' = (0 * * * *) \text{ or } \beta_{RP} = 0 \tag{2}$$

where an asterisk denotes an unrestricted parameter. As long as only bivariate linkages are examined, this is equivalent to testing the stationarity of the remaining price (PP_t) around three different levels. If the null hypotheses that $\beta_{RP} = 0$ and $\beta_{PP} = 0$ are rejected, both prices are linked in the long

10 Estimation (and hypothesis testing) using the Johansen *et al.* (2000) procedure was carried out using *GAUSS*, Version 5.0 (Aptech, 2002), and the results have been substantiated in the recently available *MALCOLM*, Version 2.5 (Mosconi, 2002), which is used with *RATS* (Estima, 2000).

run and each series is non-stationary. The second hypothesis is whether structural breaks imply changes in joint long-run price evolution; this is a test of the equality of the intercepts in the three periods and the null is

$$\mathrm{H}_0^2\colon \beta' = (*\,*\,1\,1\,1) \text{ or } \mu_1 = \mu_2 = \mu_3. \tag{3}$$

If the null is rejected, the constant component of the margin and/or the long-run equilibrium between RP_t and PP_t have not remained stable around a single level. This can be extended to test the stability of the intercept between two periods: for example, the null that the intercept is the same in the first two periods is

$$\mathrm{H}_0^3\colon \beta' = (*\,*\,1\,1\,*) \text{ or } \mu_1 = \mu_2. \tag{4}$$

4. Data and results

4.1. Data

The monthly price data for beef, lamb and pork originated from the MLC and relate to England and Wales for January 1986 to December 2000 (180 observations).[11] Producer prices are dead-weight equivalent livestock prices (pence/kg), that is, converted livestock prices using the MLC killing-out percentages of 54 per cent for beef and 45 per cent for lamb and pork. Retail prices are untrimmed average prices adjusted by drip losses of 5 per cent for beef, 3 per cent for lamb and 2 per cent for pork. The price data are illustrated in Figure 1 and summarised in Appendix 1.[12] All prices generally trend upward but in early 1996, there appears to be a structural break in each series: producer and retail beef prices fell by 6 per cent and 11 per cent, whereas corresponding prices for lamb increased both by about 20 per cent and those for pork increased by 9 per cent and 2 per cent. Since then, retail prices have generally stabilised and producer prices trend downwards. Lamb prices display seasonality.

The meat-scare index shown for descriptive purposes only in Figure 2 measures monthly references in *The Times* and *The Guardian* (and their associated Sunday publications *The Sunday Times* and *The Observer*) to BSE, *E. coli* and abattoir hygiene.[13] References to BSE constitute a large proportion of the total. The index increased substantially in early 1996 when there was increasing speculation of a causal link between BSE and Creutzfeldt–Jakob Disease (CJD); it peaked in March 1996 following the CJD Surveillance Unit identifying variant CJD in young people, and the announcement by the Secretary of State for Health that BSE was the most likely cause of their deaths (HMSO, 2000).

11 We are grateful to Sue Fisher (MLC) for providing these data.

12 Following many researchers, including Heien (1980) and von Cramon-Taubadel (1998), nominal prices were used throughout because our focus was on price behaviour across markets.

13 We are grateful to John Strak (Euro PA and Associates) for providing this index.

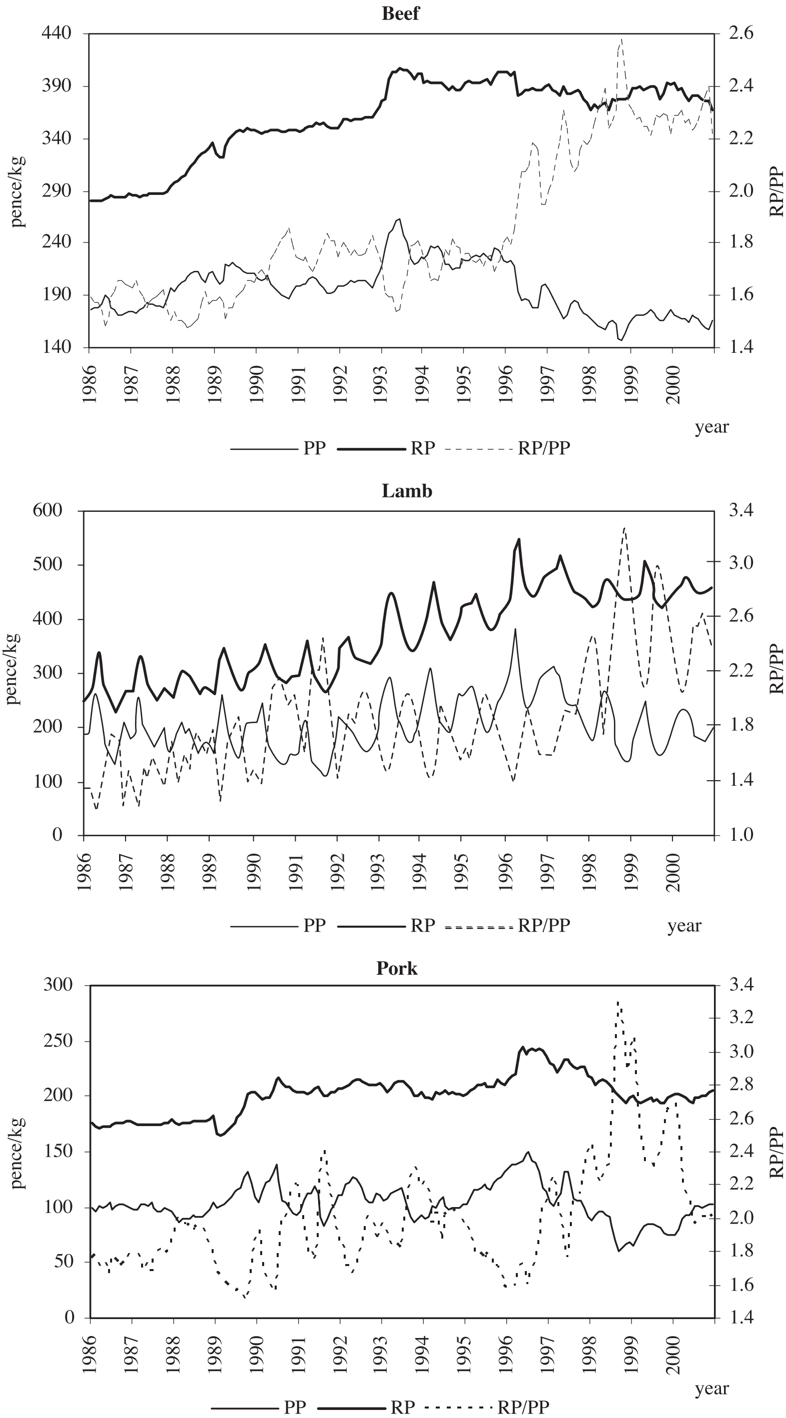

Figure 1. Producer and retail meat prices.

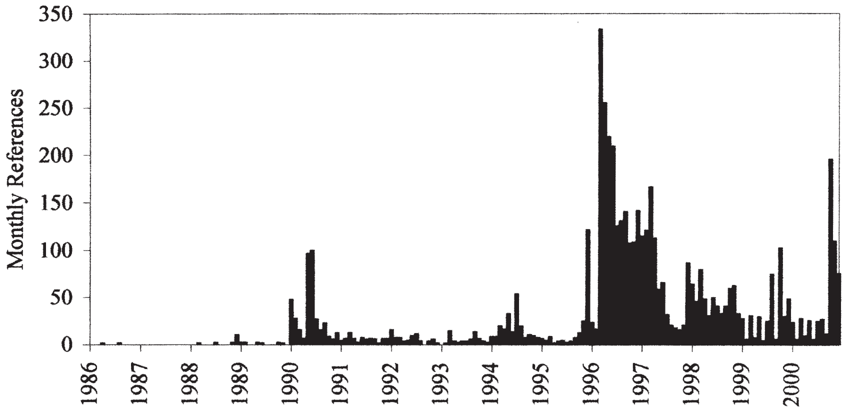

Figure 2. Meat-scare index.

4.2. Unit root tests

Perron (1997) tests were used to test for both non-stationarity and the existence of structural breaks, where the maximum lag length was 14. The results in Table 1 show that all six price series are I(1).[14] For beef, breaks occurred in September 1992 and January or February 1996; it is most likely that the break in September 1992 is related to the removal of the pound from the ERM, whereas that of early 1996 is associated with the BSE crisis. For lamb, breaks occurred in December 1992 or January 1993, February 1996 and August 1997; the last corresponds to a fall in sheep skin prices caused by reduced Russian imports following economic collapse. For pork, breaks occurred in February 1996, April 1997 and March 1998. The break in April 1997 coincides with a reduction in world supply as a consequence of pig disease outbreaks in two major producer/exporter countries (foot and mouth disease in Taiwan, and classical swine fever in the Netherlands); the consequences in the UK were higher prices and increased pig herds. The break in March 1998 coincides with large increases in production in both the UK and EU caused by strong prices in 1997. For completeness, Table 1 also shows the results for the price ratios, RP_t/PP_t. All are I(1) and breaks generally correspond to those for RP_t and PP_t. The usefulness of these results is, however, limited because the linkage between PP_t and RP_t is restricted to be contemporaneous, and dynamics are not included.

Two caveats are required concerning our use of Perron's (1997) tests. First, the determination of the break points is sensitive to the choice of lag length and different breaks result from alternative lag specifications. Nevertheless, imposing a maximum lag length of 14 is not unreasonable given the use of

14 The existence of a unit root is robust against the alternative of stationarity with two structural breaks: applying Ben-David *et al.*'s (2003) sequential approach, the minimum *t*-statistics respectively for RP_t and PP_t are: for beef, −2.60 and −5.67; for lamb, −4.49 and −3.55; and for pork, −4.90 and −5.19 (critical value at 5 per cent: −6.16).

Table 1. Perron unit root tests

Sector	Model	Price	Test statistic	Lags	Break point
Beef	Changing level	RP_t	−2.16	13	1996:2
		PP_t	−4.35	12	1996:1
		RP_t/PP_t	−4.43	12	1996:1
	Changing level and trend	RP_t	−4.45	5	1992:9
		PP_t	−4.72	12	1992:9
		RP_t/PP_t	−3.54	12	1992:9
Lamb	Changing level	RP_t	−3.38	12	1993:1
		PP_t	−3.31	12	1997:8
		RP_t/PP_t	−3.30	12	1998:5
	Changing level and trend	RP_t	−4.26	12	1996:2
		PP_t	−3.23	12	1992:12
		RP_t/PP_t	−3.15	12	1992:12
Pork	Changing level	RP_t	−3.84	12	1998:3
		PP_t	−5.00	12	1997:4
		RP_t/PP_t	−4.30	12	1997:4
	Changing level and trend	RP_t	−4.39	13	1996:2
		PP_t	−4.19	12	1997:4
		RP_t/PP_t	−4.02	12	1998:2

Critical values at the 5 per cent significance level: changing level model: −5.10; and changing level and trend model: −5.55 (Perron, 1997).

monthly data, and the number of lags chosen using a general-to-specific testing procedure are all less than this maximum. Second, the identification of two breaks in a series is problematic, as it is possible to distinguish statistically between the changing level model and the changing level and trend model, thereby leading to a preference for one against the other. In this sense, Perron's test identifies only one break point. We are unaware of similar tests with more than one break in both null and alternative hypotheses. Accordingly, we adopted a pragmatic approach by using both models to identify possible break points even if they differ. Although this method generally identifies two breaks in each series, and therefore provides an empirical inconsistency, they can, however, be identified with observed events. Moreover, we retested the unit root hypothesis with two structural breaks within the Johansen *et al.* (2000) framework.

4.3. Cointegration and hypothesis tests

The Pantula principle was used to test the joint hypothesis of both the rank of π and deterministic components in (1). For all cases, the broken level model was preferred where we allow for changes in the intercept in the cointegration space resulting from structural breaks identified in Perron's tests. This was

Table 2. Trace statistic and mis-specification tests

	Beef	Lamb			Pork		
Breaks:	1992:9 1996:2	1992:12 1996:2	1996:2 1997:8	1992:12 1997:8	1996:2 1997:4	1996:2 1998:3	1997:04 1998:03
Impulse dummies:	131	161	161	3, 4, 7, 11, 12, 16, 19, 31, 39, 40, 47, 124, 161	37, 46, 54, 67, 144, 153	37, 54, 67, 126, 137	37, 46, 54, 67, 124, 126, 144, 146
$r=0$	44.82 (32.06)	34.82 (31.93)	35.11 (30.41)	55.95 (32.03)	42.47 (30.13)	50.96 (30.71)	47.40 (29.45)
$r \leq 1$	12.26 (15.89)	8.71 (15.79)	4.46 (14.93)	5.95 (15.90)	13.93 (14.73)	16.25 (15.16)	11.93 (14.40)
Mis-specification tests							
Skewness	4.90 [0.09]	1.08 [0.58]	0.57 [0.75]	1.49 [0.47]	3.31 [0.19]	1.85 [0.39]	1.25 [0.53]
Kurtosis	2.66 [0.26]	2.66 [0.26]	2.82 [0.24]	7.09 [0.03]	5.47 [0.06]	9.09 [0.01]	11.08 [0.00]
Skewness and kurtosis	7.56 [0.11]	3.75 [0.44]	3.39 [0.49]	8.59 [0.07]	8.78 [0.07]	10.95 [0.03]	12.33 [0.01]
Autocorrelation (Portmanteau test)	37.42 [0.59]	55.03 [0.06]	54.36 [0.06]	45.82 [0.24]	51.76 [0.10]	45.64 [0.25]	45.00 [0.27]
ARCH(2)							
RP_t	0.71 [0.70]	2.13 [0.34]	3.90 [0.14]	4.84 [0.09]	3.74 [0.15]	9.27 [0.01]	5.21 [0.07]
PP_t	0.13 [0.94]	0.46 [0.79]	2.44 [0.29]	0.23 [0.89]	1.82 [0.40]	3.31 [0.19]	3.36 [0.19]

Impulse dummies denoted at each observation. Critical values in parentheses at the 5 per cent significance level. *p*-values in square brackets.

expected: there are trends in the levels of the series but not in differences so that each VECM includes a constant in the long run (Franses, 2001), which is the constant component of the margin.

Each model contains two sets of dummy variables. First, 11 centred seasonal dummies were included to account for deterministic seasonality. Second, following common practice, impulse dummies were included to account for outliers to produce well-specified residuals, u_t, and they were included if the corresponding residual is more than twice the standard error. The Schwarz criterion (Lütkepohl, 1993: 132) was used to determine the lags in each VECM, and in all cases $p = 2$. Table 2 reports trace statistics to test for the number of cointegrating vectors (r),[15] mis-specification tests of the residuals—namely, multivariate normality statistics for skewness and kurtosis,

15 Harris (1995: 81) noted that the distribution of the rank test might be influenced by the inclusion of short-run impulse dummies and Appendix 2 presents results with seasonal dummies only.

Table 3. Likelihood ratio tests on the long-run parameters

Hypothesis	Beef	Lamb			Pork		
	1992:9 1996:2	1992:12 1996:2	1996:2 1997:8	1992:12 1997:8	1996:2 1997:4	1996:2 1998:3	1997:04 1998:03
H_0^1: $\beta_{RP} = 0$	6.84 [0.00]	5.90 [0.01]	10.54 [0.00]	3.81 [0.05]	7.50 [0.00]	11.05 [0.00]	7.86 [0.00]
H_0^1: $\beta_{PP} = 0$	19.55 [0.00]	5.98 [0.01]	26.04 [0.00]	39.24 [0.00]	14.39 [0.00]	18.29 [0.00]	23.54 [0.00]
H_0^2: $\mu_1 = \mu_2 = \mu_3$	17.20 [0.00]	12.21 [0.00]	17.00 [0.00]	32.20 [0.00]	8.53 [0.01]	8.81 [0.01]	12.65 [0.00]
H_0^3: $\mu_1 = \mu_2$	2.65 [0.10]	11.43 [0.00]	1.98 [0.16]	8.27 [0.00]	0.01 [0.90]	1.51 [0.22]	7.46 [0.00]
H_0^3: $\mu_2 = \mu_3$	16.23 [0.00]	1.43 [0.23]	15.95 [0.00]	23.85 [0.00]	3.69 [0.05]	1.42 [0.23]	0.51 [0.48]

p-values in square brackets using χ_1^2 for H_0^1 and H_0^3, and χ_2^2 for H_0^2.

portmanteau tests for autocorrelation up to order 12 (Lütkepohl, 1993: 150–158), and autoregressive conditional heteroscedasticity (ARCH) tests of order 2 (Johnston and NiDardo, 1997: 195–196). Table 3 reports the results of the hypothesis tests.

In the beef model, two breaks in September 1992 and February 1996 are included.[16] Trace statistics show that $r = 1$ and a long-run relationship exists between PP_t and RP_t; mis-specification tests show that the model is well specified. Regarding the hypotheses of interest, we first tested the nulls (H_0^1) that each price belongs to the cointegration space. Both hypotheses were rejected, implying that PP_t and RP_t are linked in the long run, and that each is non-stationary, which substantiates the conclusions from Perron's tests.[17] Second, the null (H_0^2) that the intercepts in the three periods are equal was rejected, implying that the constant component of the margin and long-run equilibrium between RP_t and PP_t have not remained stable. Third, we tested the stability of the intercept before and after breaks (H_0^3). The null hypothesis that $\mu_2 = \mu_3$, where μ_2 is the constant component of the margin between September 1992 and February 1996 and μ_3 is the constant after the 1996 BSE events, was rejected. Thus, the BSE crisis had a significant impact on long-run equilibrium, affecting both prices individually and altering the relationship that keeps them together. The hypothesis that $\mu_1 = \mu_2$, where μ_1 is the constant component before September 1992, was not rejected and the breaks in both RP_t and PP_t cancel each other out. Imposing this

16 The results from a model with a break in January 1996 are similar.

17 Note the tension between the two approaches: in Perron's test, only one endogenously estimated break is allowed, whereas in the Johansen *et al*. procedure up to two predetermined breaks are permitted.

restriction ($\mu_1 = \mu_2$), the cointegration vector is

$$RP_t = 2.36PP_t - 130.56E_1 - 130.56E_2 - 18.72E_3. \qquad (5)$$

These estimates imply that there is a percentage mark-up of 236% but, counter-intuitively, the absolute component of the margin is negative, being £1.31/kg before the BSE crisis and £0.19/kg thereafter. Nevertheless, the margin calculated from the estimates in (5) is always positive. Recognising Gardner's caveat that our simple bivariate price relationship is an approximation to the true relationship, we focus on the relative magnitudes of the constants in (5). Thus, the increase in the margin following the BSE crisis was £1.12/kg, which accords with Gardner's theory.

For lamb, three models were estimated with breaks in December 1992 and February 1996, in February 1996 and August 1997, and in December 1992 and August 1997. Trace statistics show that $r = 1$ in all models; all models are well specified although there is some evidence of mild excess kurtosis in that with breaks in December 1992 and August 1997.[18] Hypothesis tests (H_0^3) imply that the BSE crisis had no impact on the long-run equilibrium; accordingly, the model with breaks in December 1992 and August 1997, which are significant, was preferred. Using this model, the nulls H_0^1 were rejected: each price belongs to the cointegration space, both prices are linked in the long run, and each series is non-stationary, again substantiating the conclusions from Perron's tests. The cointegrating vector is

$$RP_t = 3.81PP_t - 344.16E_1 - 515.85E_2 - 272.27E_3. \qquad (6)$$

With the same caveat as for the beef model in (5), the shock in December 1992 had a permanent effect on the long-run equilibrium, which reduced the margin by £1.72/kg and was in line with expectations. Support for sheepmeat was largely in the form of various ewe subsidies, which were set in ECUs. The removal of the pound from the ERM, and the subsequent devaluation, resulted in increases in these subsidies in pounds sterling, and a lagged increase in the producer price of lamb. The retail price of lamb also increased after a lag following the exchange rate depreciation. This occurred partly because exports increased, which caused supplies to the domestic market to fall, and partly because the price of imports rose.[19] From Gardner's theory, both these effects lead to a reduction in the margin. The shock in August 1997 following the fall in sheep skin prices increased the margin by £2.44/kg. This accords with Gardner's theory if the fall in demand for sheep skins reduces demand for the joint product (sheepmeat/skins).

18 In terms of testing for cointegration, Cheung and Lai (1993) show that 'The trace test . . . shows little bias in the presence of either skewness or excess kurtosis.' In terms of estimation, Gonzalo (1994) shows that 'Johansen's procedure performs better than [other cointegration] methods even when the errors are nonnormally distributed'.

19 In 1993, sheepmeat exports increased by 22 per cent whereas supplies to the UK market fell by 11 per cent (HMSO, 1994: 41). Between December 1992 and April 1993, deseasonalised RP_t and PP_t increased by 21 per cent and 36 per cent whereas the average changes over the previous 5 years were +3 per cent and −6 per cent.

For pork, three models were estimated with breaks in February 1996 and April 1997, in February 1996 and March 1998, and in April 1997 and March 1998. Trace statistics show that $r = 1$ in all models, and each is generally well specified although there is evidence of excess kurtosis in the latter two and heteroscedasticity in the second. Hypothesis tests (H_0^3) imply that the BSE crisis and the increases in production in 1998 had no impact on long-run equilibrium. Because our focus was on the BSE crisis, our preferred model was with breaks in February 1996 and April 1997 where the latter is significant. Using this model, the nulls H_0^1 were rejected: each price belongs to the cointegration space, both prices are linked in the long run, and each series is non-stationary, again substantiating the conclusions from Perron's tests. The cointegrating vector is

$$RP_t = 1.80PP_t + 6.94E_1 + 6.94E_2 + 39.82E_3. \tag{7}$$

Comparing the intercepts, the increase in the margin resulting from reduced world supply in April 1997 was £0.33/kg. Assuming this reduction in supply increases the retail price of pork, this effect accords with Gardner's theory.

To emphasise the importance of adequately modelling the breaks, an additional model was estimated for each sector with no breaks. Both models were used to simulate values of RP_t/PP_t after the first significant break; that is, after February 1996 for beef, December 1992 for lamb, and April 1997 for pork. We substituted lagged values of the variables in both models by predictions to compare the accuracy of the projections, and the results are shown in Figure 3. For each sector, but particularly for beef, our preferred model with breaks provides a more accurate forecast than the model without.[20] Furthermore, RP_t/PP_t for beef increased substantially in the period after the BSE crisis in early 1996, which accords with Gardner's theoretical predictions.

These results require two caveats. First, the estimated parameters in the cointegrating vectors (β) are sensitive to the choice of break points and alternatives can produce dramatic shifts in their coefficients. This occurred for lamb and pork, whereas the beef model was robust to specification changes. Second, the Perron tests identified three breaks in the prices of lamb and pork that correspond to observable events in their respective markets; as Johansen *et al.*'s procedure permits up to two breaks only, we were required to choose between models with alternative pairs of breaks.

5. Conclusions

From the late 1980s, public concern increased in the UK about the effects on human health of BSE transmission from cattle. The critical period was early 1996 when the Government announced that a death from variant CJD was

20 For beef, the percentage root-mean-square error 6 months ahead is 2 per cent for the model with breaks and 9 per cent for the model without breaks. Corresponding values for lamb are 5 per cent and 11 per cent, and those for pork are 5 per cent and 8 per cent.

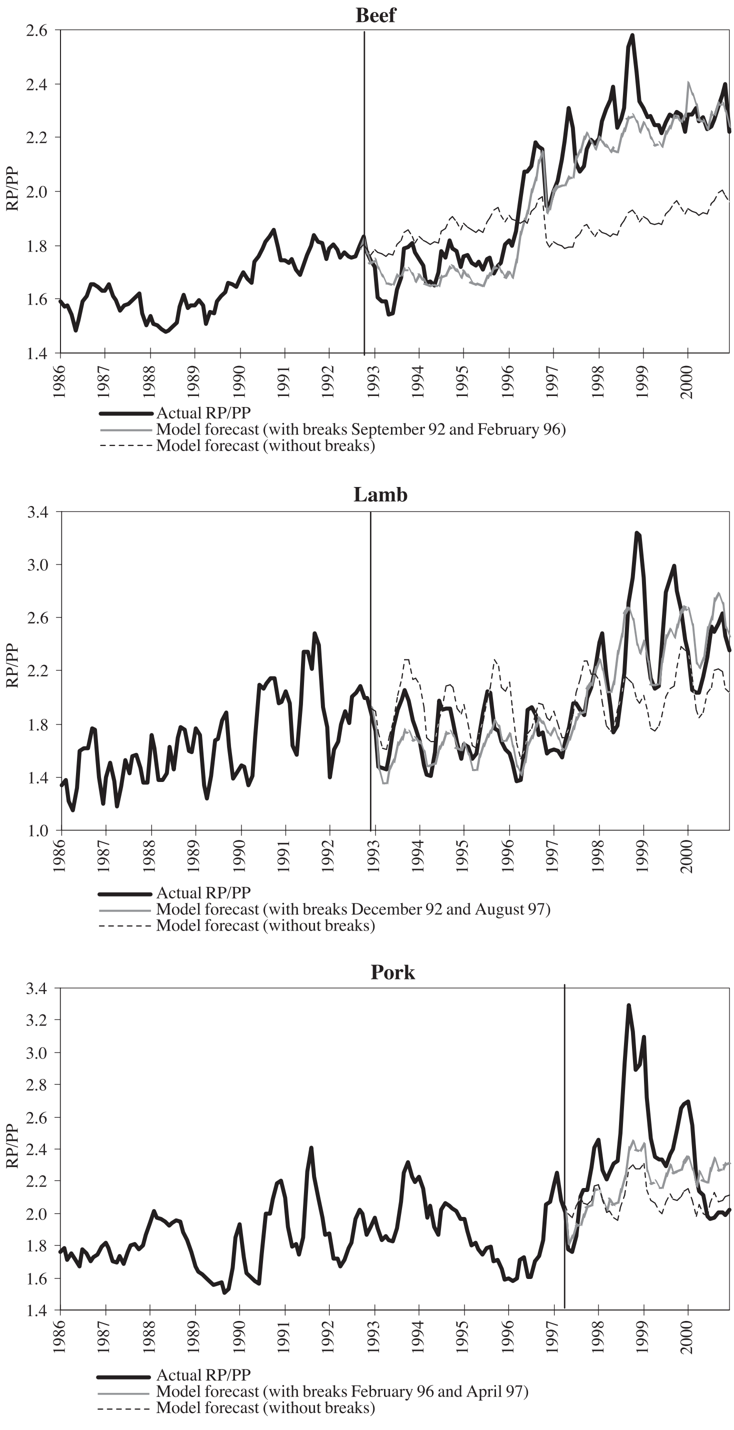

Figure 3. Actual and forecast price ratios.

probably linked to BSE. This paper examines producer–retail price transmission for beef, lamb and pork and the impact of the BSE crisis. Using monthly data for England and Wales for 1986–2000, we have sought long-run relationships between producer and retail prices for each meat using the cointegrating procedure of Johansen *et al.* (2000), which allows for up to two structural breaks.

For each meat, results show that a long-run relationship exists between producer and retail prices, and structural breaks occurred in February 1996 for beef, in December 1992 and August 1998 for lamb, and in April 1997 for pork. In the lamb relationship, the cause of the break in 1992, which resulted in a fall in the margin of £1.72/kg, is not clear: although CAP reform was taking place at the time, the more likely cause is a lagged response from the removal of the pound from the ERM on Black Wednesday (16 September). The break in August 1998 corresponds to a collapse of Russian sheep skin imports and resulted in a rise in the margin of £2.44/kg. In the pork relationship, the break in 1997, when the margin increased by £0.33/kg, was probably caused by reduced world supply following foot and mouth disease in Taiwan and swine fever in the Netherlands.

Structural breaks occurred in September 1992 and February 1996 in both retail and producer prices of beef. The breaks in September 1992 cancelled each other out and there was no impact on the margin. In contrast, the breaks in February 1996, which were caused by the dissemination in the media of information about BSE, did not cancel each other out and there was a structural break in the beef-price relationship; the producer–retail margin increased by £1.12/kg and consumer and farmer perceptions that middlemen and supermarkets benefited from the crisis appear to be substantiated. Under the plausible assumption that agricultural commodities are more inelastic than marketing services, these results accord with Gardner's (1975) theoretical propositions. The BSE crisis seems not to have affected price transmission in either the lamb or pork sectors.

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Appendix 1: Description of data, January 1986–December 2000 (pence/kg)

Sector	Price	Mean	SD	Minimum	Maximum
Beef	Retail	358.94	37.44	279.98	407.71
	Producer	196.16	24.21	146.15	263.70
Lamb	Retail	366.70	78.26	225.97	542.36
	Producer	204.12	49.01	108.94	379.51
Pork	Retail	201.15	18.23	164.94	245.30
	Producer	103.96	16.99	60.91	150.75

In December 2000, 1€ = £0.6134.

Appendix 2: Trace statistics and mis-specification tests for models that include a full set of seasonal dummies

	Beef	Lamb			Pork		
Breaks:	1992:9 1996:2	1992:12 1996:2	1996:2 1997:8	1992:12 1997:8	1996:2 1997:4	1996:2 1998:3	1997:4 1998:3
$r = 0$	37.76 (32.06)	34.94 (31.93)	34.70 (30.41)	45.41 (32.03)	30.22 (30.13)	27.56 (30.71)	34.84 (29.45)
$r \leq 1$	11.61 (15.89)	7.79 (15.79)	4.45 (14.93)	8.79 (15.90)	8.59 (14.73)	6.63 (15.16)	7.05 (14.40)
Mis-specification tests							
Skewness	10.83 [0.00]	1.73 [0.42]	1.02 [0.60]	32.93 [0.00]	56.43 [0.00]	42.86 [0.00]	2.90 [0.23]
Kurtosis	8.66 [0.01]	0.20 [0.90]	1.09 [0.58]	196.71 [0.00]	383.67 [0.00]	317.90 [0.00]	449.05 [0.00]
Skewness and kurtosis	19.50 [0.00]	1.93 [0.75]	2.12 [0.71]	229.65 [0.00]	440.09 [0.00]	360.76 [0.00]	451.95 [0.00]
Autocorrelation (Portmanteau test)	33.86 [0.74]	65.97 [0.01]	62.26 [0.01]	56.89 [0.04]	51.54 [0.10]	47.57 [0.19]	51.07 [0.11]
ARCH(2)							
RP_t	0.82 [0.66]	3.41 [0.18]	2.90 [0.23]	0.58 [0.75]	0.05 [0.98]	0.07 [0.97]	1.22 [0.54]
PP_t	0.13 [0.93]	0.50 [0.78]	1.64 [0.44]	1.47 [0.48]	12.42 [0.00]	8.57 [0.01]	10.82 [0.00]

Critical values in parentheses at the 5 per cent significance level. *p* values in square brackets.

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