

Market Power and Relative Price Adjustment: Evidence from the UK

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Abstract

Empirical studies of price transmission often suggest that imperfect pass-through may be due to market power exerted by food retailers. However, these econometric studies essentially lack any formal basis for tying the role of market power with data comprising of retail and producer prices only. We show that if market power has an effect on the farm-retail margin, this determines the specification of the cointegrating relationship. To emphasise the relevance of the tests, we focus on results relating the UK beef sector and show that market power is likely to have played a role in determining the retail-farm price margin.

Keywords: price adjustment; market power

JEL Classification: Q11; L13

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1. Introduction

In recent years, there have been a large number of empirical studies that have focused on price transmission between the retail and farm sectors with a view to testing whether price changes at one stage are fully passed through to the other. Typically, in the absence of perfect price transmission, the authors will assume that one important candidate for imperfect price transmission is market power in the stages linking the retail and farm sectors. A typical (though simplified) example of this would be to have data on price pairs (retail and producer), test whether there is evidence of cointegration and explore the dynamics linking these prices together. In the absence of a cointegrating relationship and/or the result of perfect (often unitary) price transmission between prices, there is an appeal to the theoretical literature on price transmission which highlights the role of market power (see below). However, while the econometric techniques applied can have varying degrees of sophistication, little effort is paid to the underlying structure of the basic model, particularly if one is trying to identify a role for market power in determining the result. This is the issue that is addressed in this paper. Specifically, we outline a procedure that will determine the structure of the cointegrating regression if market power is likely to be present. The attractive feature of this test is that it is (a) essentially simple to apply and (b) provides a more formal basis for addressing the role of market power on price transmission in the food chain and (c) avoids the researcher appealing to ad hoc factors in explaining imperfect price transmission that several econometric studies typically find.

It should also be noted that the issue of retail-farm level price transmission is not confined to academic debate but is also frequently highlighted in public policy as a means for inferring the misuse of market power. For example, several studies by the UK anti-trust authority (the Competition Commission and its predecessor the Monopolies and Mergers Commission) have often highlighted the role of market power in determining how price relationships should evolve. To take the recent report into market power potentially exercised by food retailers, emphasis was given to price transmission both in terms of the reasons for the investigation in the first place and in terms of evidence based on econometric studies relating to the dynamics of price adjustment at the retail and farm levels¹. A more recent report commissioned by the UK Department for the Environment, Food and Rural Affairs (DEFRA, 2004) also focused on price transmission as a follow up to the Competition Commission report with the remit being whether market power potentially exercised by retailers influenced the price transmission relationship. The latter report highlights the type of issues that this paper addresses since a close reading of this report casts doubt on whether their results can really substantiate the claims being made. This is because there is little tie in with the theory with the empirics and therefore makes their econometric strategy and the subsequent interpretation of their results questionable.

¹ The Competition Commission noted one of their main concerns being to address: '...[the] public perception of...an apparent disparity between farm-gate and retail prices...which is seen as evidence by some that grocer multiples were profiting from the crisis in the farming industry'. (Competition Commission, Vol. 1, p.3, 2000)

The paper is organised as follows. In section 2, we briefly review the theoretical literature linking price transmission with market power in the food sector. In section 3, we outline a simple theoretical model that forms the basis for the structure of the cointegrating relationship. This framework forms the basis for our empirical tests which we apply to 10 commodity sectors using retail and farm prices in the UK. However, given space constraints, we focus solely on the results from the UK beef sector in some detail. This is reported in section 4. We also outline briefly the results for other commodity sectors. Section 5 summarises and concludes.

2. Linking Price Transmission with Market Power

There is a broad literature on the issue of the margin between the retail and farm levels and what factors may influence it. The most notable early paper on this issue was by Gardner (1975) which identified a range of factors that would influence the price transmission between the farm and retail sectors. Gardner assumed perfect competition which clearly does not fit with the concerns raised by the UK anti-trust authorities. To this end, McCorrison *et al.* (1998) show that oligopoly power in the food sector would have an impact on determining the price transmission elasticity following a supply side shock depending on the functional form of the demand curve while McCorrison *et al.* (2001) show that the extent of returns to scale characterising the food industry cost function will also be important. Other important influences of the retail-farm margin and hence price transmission are likely to be oligopsony power (Lloyd *et al.*, 2002), and the source of the exogenous shock (i.e. whether the shift occurs in the retail demand or farm supply function (Gardner, *op. cit.*)).

In the theoretical framework outlined below, we draw on the previous literature to highlight the potential role of market power on the retail-farm margin. However, rather than derive an explicit price transmission elasticity, we use the framework to determine the econometric strategy in linking price transmission with market power. Specifically, we show that if market power characterises the UK food sector then both the exogenous demand and supply shifters should enter the reduced form retail-farm margin equation. If market power is not a feature of the sector, then there would be no *a priori* case for the inclusion of the shifters and the margin simply reflects marketing costs. Therefore, while we do not retrieve an explicit measure of market power, market power will nevertheless influence the outcome if the demand and supply shifters are found to be statistically significant in the reduced form model of the retail-farm gate margin. In addition, we also relate the results to the derivation of impulse response functions since these should be explicitly about the impact of a shifter on relative prices rather than a change in one of the prices only which therefore ties in precisely with the theoretical work on price transmission.

3. Theoretical Framework

The demand function for the processed product is given by:

$$Q = h(R, R^s, X) \quad (1)$$

where R is the retail price of the good under consideration and R^s is the price of a substitute good which firms in this sector take as given. X is the demand shifter. The supply function of the agricultural raw material is given by (in inverse form):

$$P = k(A, N) \quad (2)$$

where A is the quantity of the agricultural raw material and N is the exogenous shifter in the farm supply equation.

For a representative firm, the profit function is given by:

$$\pi_i = R(Q)Q_i - P(A)A_i - C_i(Q_i) \quad (3)$$

where C_i is other costs and, assuming a fixed proportions technology, $Q_i = A_i / a$ where a is the input:output coefficient which is assumed to equal 1. This assumption corresponds closely to the construction of the data in the vertical market chain used in the empirical analysis that follows². The first-order condition for profit maximisation is given by:

$$R + Q_i \frac{\partial R}{\partial Q} \frac{\partial Q}{\partial Q_i} = \frac{\partial C_i}{\partial Q_i} + aP + aA_i \frac{\partial P}{\partial A} \frac{\partial A}{\partial A_i} \quad (4)$$

In order to get an explicit solution, consider linear functional forms for equations (1) and (2) and assume $a = 1$:

$$Q = h - bR + eR^s + cX \quad (1')$$

$$P = k + gS \quad (2')$$

with domestic supply being given by:

$$S = Q + N$$

where N is the level of exports which are exogenously determined. From this, and a aggregating over n -firms, (4) can be re-written as:

$$R - \frac{\theta}{nb}Q = M + P + \frac{\mu g Q}{n} \quad (4')$$

where θ and μ represent the conjectures relating to oligopoly and oligopsony power respectively. These parameters can be interpreted as an index of market power with $\theta = \mu = 0$ representing competitive behaviour and $\theta = \mu = 1$ representing collusive behaviour. M represents other costs that enter the industry cost function which are assumed to be marketing costs, the price of which is taken as given. Assume for ease of interpretation, θ and μ are n -firm weighted indices of market power where n is small. Using (1'), (2') and (4'), we can derive an explicit solution for the endogenous variables:

$$Q = \frac{h + cX + eR^s - b(M + k + gN)}{(1 + \theta) + bg(1 + \mu)} \quad (5)$$

$$R = \frac{\theta(h + cX + eR^s) / b + g(1 + \mu)(h + cX + eR^s) + M + k + gN}{(1 + \theta) + bg(1 + \mu)} \quad (6)$$

$$P = \frac{(1 + \theta)(k + gN) + bg\mu(k + gN) + g(h + eR^s + cX - bM)}{(1 + \theta) + bg(1 + \mu)} \quad (7)$$

To derive the retail-farm spread, use (6) and (7) to give:

² In the commodity sectors studied, care was taken to select products where the technology would be most obviously characterised as fixed proportions.

$$R - P = \frac{M(1 + bg) - (\theta + bg\mu)(k + gN) + (\theta/b + g\mu)(h + eR^s + cX)}{(1 + \theta) + bg(1 + \mu)} \quad (8)$$

Note that if neither oligopoly nor oligopsony power matters in the determining the retail-farm price spread (i.e. $\theta = \mu = 0$), then equation (8) reduces to:

$$R - P = M \quad (9)$$

i.e. the source of the retail-farm price margin in a perfectly competitive industry is due to marketing costs only. In this case, the role of the exogenous shifters play no role in determining the spread. This is not to say that they do not affect each price individually, but in a perfectly competitive industry they play no role in determining the relative gap between the prices at each stage of the food chain. Correspondingly, if either oligopoly and/or oligospony power in the food sector is important, then they will influence the margin between retail and farm prices. In other words, each shifter will affect the two prices differentially and thus the margin between the prices will change.

Equations (6)-(8) form the basis for the structure of the econometric model. Consider first of all equation (8) that relates to the retail-farm spread. Note that if market power in some form does characterise the UK food sector, then the exogenous supply and demand shifters should enter our econometric model of the margin between retail and farm prices. Hence there are two aspects to linking theory on market power and price transmission and the specification of the econometric model. First, one should test for cointegration in the presence of exogenous shifters rather than price pairs alone. The latter is insufficient to draw any implications about the potential role of market power in price relationships. Second, and following from the specification of the cointegrating relationship, the test for the existence of market power is whether the coefficients on these variables in the retail-farm spread equation are statistically significant. If market power does play a role, then this will influence the retail and farm level prices to varying degrees. Finally, to the extent that (either of) these shifters are significant, we can therefore use (6) or (7) with (8) to derive the impact of exogenous shifters on retail and farm prices with N capturing the farm-level shifts and X capturing the impact of the demand shifter at the retail level using impulse response functions to trace out the implied price transmission relationship between farm and retail prices. Note however that the source of shock that underpins the impulse response relates to the presence of the shifter at either the farm or retail level (or both) and hence directly links the theoretical model with the econometric approach and not a ‘shock’ characterised by a change in one of the prices that forms the price pair.

We collected data for 10 commodity sectors in the UK and cover the period 1990-2002. The data covers farm and retail prices for the following sectors: beef, lamb, pork, chicken, sugar, milk, fresh fruit, coffee and eggs. To characterise the shifters, we separated the meats sectors from others in that over the data period, the meats sector was affected by the BSE crisis in the mid-1990s. In this case, we had data that explicitly characterised the food scare and the impact of the export ban and the cull in the beef sector. For the other sectors, the demand shifter related to the UK food retail price index on the basis that any shocks to demand will be indicated by changes in

this overall index. To capture supply shocks, we used the index of prices of purchases by UK agriculture on the basis that any shocks to agriculture will be reflected in the derived demand for farm inputs. However, due to space constraints and the technicalities associated with the econometric tests, for present purposes we focus only on the specification for the UK beef sector. Specifically, in the following empirical section, we apply this theoretical framework using data from the UK beef market that explicitly accounts for demand and supply side shifters.

4. Econometric Issues

General

In this section, we detail the econometric methodology relating to issues identifying cointegration in the presence of shifters through to the use of impulse response functions where the shifters themselves may be endogenous in the long-run relationship between all variables in the system. As a starting point note that since prices are likely to be non-stationary and cointegrated, it is appropriate to couch the empirical analysis in a vector autoregressive (VAR) framework (Hendry and Doornik, p129, 2001). Consider a VAR(p) model:

$$\mathbf{x}_t = \Phi_1 \mathbf{x}_{t-1} + \Phi_2 \mathbf{x}_{t-2} + \dots + \Phi_p \mathbf{x}_{t-p} + \Psi \mathbf{w}_t + \boldsymbol{\varepsilon}_t \quad (10)$$

where \mathbf{x}_t is a ($m \times 1$) vector ($1, 2, \dots, i, j, \dots, m$) of jointly determined I(1) variables, \mathbf{w}_t is a ($q \times 1$) vector of deterministic and or exogenous variables and each Φ_i ($i = 1, \dots, p$) and Ψ are ($m \times m$) and ($m \times q$) matrices of coefficients to be estimated by Johansen's (1988) maximum likelihood procedure using a ($t = 1, \dots, T$) sample of data. $\boldsymbol{\varepsilon}_t$ is a ($m \times 1$) vector of n.i.d. disturbances with zero mean and non-diagonal covariance matrix, Σ .

The error correction representation of (10) is observationally equivalent but facilitates estimation and hypothesis testing since all terms are stationary (Hendry and Doornik, p.60, *op. cit.*). This re-parameterisation is given by:

$$\Delta \mathbf{x}_t = \boldsymbol{\alpha} \boldsymbol{\beta}' \mathbf{x}_{t-p} + \sum_{i=1}^{p-1} \Gamma_i \Delta \mathbf{x}_{t-i} + \Psi \mathbf{w}_t + \boldsymbol{\varepsilon}_t \quad (11)$$

Attention focuses on the ($n \times r$) matrix of co-integrating vectors, $\boldsymbol{\beta}$, that quantify the 'long-run' (or equilibrium) relationships between the variables in the system and the ($n \times r$) matrix of error correction coefficients, $\boldsymbol{\alpha}$, the elements of which load deviations from equilibrium (*i.e.* $\boldsymbol{\beta}' \mathbf{x}_{t-k}$) into $\Delta \mathbf{x}_t$, for correction. The Γ_i coefficients in (11) estimate the short-run effect of shocks on $\Delta \mathbf{x}_t$, and thereby allow the short and long-run responses to differ.

Given the proceeding discussion, $\boldsymbol{\beta}$ may represent the economic linkages that bind the prices together in the long run. As section 3 demonstrates, these linkages occur either between substitutes at the retail level or between marketing stages for a single good and as such provide the identifying restrictions required to interpret the cointegrating relationships in an economically meaningful way.³ However, as Lütkepohl and Riemers (1992) make clear, despite offering estimates of the 'long

³ Like any set of equations, the usual rank and order conditions must be satisfied if $\boldsymbol{\beta}$ is to be unique and the cointegration relationships given valid economic interpretations.

run', they are by construction partial derivatives predicated on the *ceteris paribus* assumption. When the variables in a cointegrated system are characterised by rich dynamic interaction (such that x_1 affects x_2 and x_2 affects x_1 , possibly with lags and/or through other variables), inference based upon 'everything else held constant' may have limited value. If what is actually wanted is an estimate of what might happen following a specific shock, then impulse response analysis, which takes account of these interactions, provides a tractable and potentially attractive solution. Whilst impulse response functions are readily calculated from reduced forms such as (11), knowledge of the structural economic representation from which (11) derives is required if the correlations upon which the impulse response functions are based are to be legitimately attributed to economic (causal) mechanisms.

As we shall see, it is the dynamic effect of shocks to the supply and demand shifters that are sought. While these have an intuitively 'structural' interpretation, in that the shifters reflect influences that can be reasonably thought of as driving prices rather than being driven by them, it is straightforward to formalise. To do so consider the structural economic representation of (11), namely:⁴

$$\mathbf{A} \Delta \mathbf{x}_t = \tilde{\boldsymbol{\alpha}} [\boldsymbol{\beta} \mathbf{x}_{t-p}] + \sum_{i=1}^{p-1} \tilde{\boldsymbol{\Gamma}}_i \Delta \mathbf{x}_{t-i} + \tilde{\boldsymbol{\Psi}} \mathbf{w}_t + \mathbf{v}_t \quad (12)$$

where \mathbf{A} represents the 6×6 matrix of coefficients defining the contemporaneous structural linkages in the system, $\tilde{\boldsymbol{\alpha}} = \mathbf{A} \boldsymbol{\alpha}$, $\tilde{\boldsymbol{\Gamma}}_i = \mathbf{A} \boldsymbol{\Gamma}_i$, $\tilde{\boldsymbol{\Psi}} = \mathbf{A} \boldsymbol{\Psi}$ and

$$\mathbf{v}_t = \mathbf{A} \boldsymbol{\varepsilon}_t \quad (13)$$

are the structural shocks which, as pure disturbances, are assumed to be serially uncorrelated and uncorrelated with each other with zero means and diagonal variance-covariance matrix, $\boldsymbol{\Omega} = E[\mathbf{v}_t \mathbf{v}_t'] = \mathbf{A} \boldsymbol{\Sigma} \mathbf{A}'$ (Hamilton, 1994, p.329). The 6^2 restrictions on \mathbf{A} (one normalisation and five homogenous linear restrictions per equation) required for exact identification of (12) is overly rigorous for our purposes since all we require is that the two shifter equations be identified. Decomposing $\mathbf{x}_t = (\mathbf{x}_{1t}, \mathbf{x}_{2t}')'$ where, using the notation from the previous section, $\mathbf{x}_{1t} = (X_t, N_t)$ denotes the shifters and $\mathbf{x}_{2t} = (R_t, R_t^s, P_t)$ denotes the prices, we may partition the structural model (12) accordingly, to yield:

$$\begin{pmatrix} \mathbf{A}_{11} & \mathbf{A}_{12} \\ \mathbf{A}_{21} & \mathbf{A}_{22} \end{pmatrix} \begin{pmatrix} \Delta \mathbf{x}_{1t} \\ \Delta \mathbf{x}_{2t} \end{pmatrix} = \boldsymbol{\mu}_{t-1} + \begin{pmatrix} \mathbf{v}_{1t} \\ \mathbf{v}_{2t} \end{pmatrix} \quad (14)$$

where

$$\boldsymbol{\mu}_{t-1} = \tilde{\boldsymbol{\alpha}} [\boldsymbol{\beta} \mathbf{x}_{t-p}] + \sum_{i=1}^{p-1} \tilde{\boldsymbol{\Gamma}}_i \Delta \mathbf{x}_{t-i} + \tilde{\boldsymbol{\Psi}} \mathbf{w}_t$$

and

$$\begin{pmatrix} \mathbf{v}_{1t} \\ \mathbf{v}_{2t} \end{pmatrix} \sim n.i.d \left[\mathbf{0}, \begin{pmatrix} \boldsymbol{\Omega}_{11} & \boldsymbol{\Omega}_{12} \\ \boldsymbol{\Omega}_{21} & \boldsymbol{\Omega}_{22} \end{pmatrix} \right]$$

⁴ Shocks to (say) the price of a substitute are less obviously 'structural' since they themselves are likely to reflect shocks to other meat prices.

Exact identification of the equations in $\Delta \mathbf{x}_{1t}$ (and the impulse response functions associated with them) is achieved providing:

$$\mathbf{A}_{12} = \begin{pmatrix} 0 & 0 \\ 0 & 0 \end{pmatrix}, \quad \mathbf{A}_{11} = \begin{pmatrix} 1 & 0 \\ 0 & 1 \end{pmatrix} \quad \text{and} \quad \mathbf{\Omega}_{11} = \begin{pmatrix} \omega_X & 0 \\ 0 & \omega_N \end{pmatrix}.$$

The first set of restrictions state that prices are not contemporaneously causal for the shifters; the second, that the shifters themselves are contemporaneously unrelated; and the third, that shocks to the shifters are orthogonal to each other. Given that we can safely assume that the contemporaneous causality is from the shifters to prices rather than *vice versa* and that each shifter reflects changes in disparate parts of the economy and are thus unrelated, the equations comprising $\Delta \mathbf{x}_{1t}$ are exactly identified. This is helpful empirically since the impulse response functions of the variables contained in $\Delta \mathbf{x}_{1t}$ do not require the equations in $\Delta \mathbf{x}_{2t}$ to be identified nor is $\mathbf{\Omega}_{12} = 0$ a requirement so that shocks between prices and shifters need not be orthogonal.⁵

In what follows, we use the generalised impulse response function (Koop *et al.* 1996, Pesaran and Shin, 1998) to assess the impact of demand and supply side shocks that characterised the BSE crisis in the UK on beef prices at both retail and farm levels. These are used to: (a) infer whether the food sector is characterised by market power as reflected in equation (8) and, (b) gauge the relative importance of the demand and supply shocks that characterise the BSE crisis arising by combining equation (6) or (7) with (8). Note then there is an explicit tie with market power and price transmission and the role of the shifters in relation to the impulse response functions which is consistent with the theoretical approach to this issue.

Data for the Beef Sector

Monthly price data spanning January 1990 to December 2000 are supplied by the UK's Department of Environment, Food and Rural Affairs (DEFRA).⁶ Retail price series on beef, pork, and lamb are those collected as part of the UK's Meat and Livestock Commission Retail Prices Survey. The survey covers purchases in a variety of retailers (such as independent butchers and supermarkets) in 21 locations in England and Wales. A representative retail price for each meat is constructed through aggregation of prices recorded for individual cuts according to their share in a carcass. The producer price of beef is derived from a weekly survey of average live-weight prices at 190 auctions market in Great Britain. All prices have been deflated by the retail price index (December 1999 base) and are measured in pence per kilogram (p/kg). To facilitate comparison between retail and producer levels of the marketing chain all prices are expressed in 'carcass weight equivalents'.

These price data are augmented by two variables representing demand and supply shifters in the UK beef market. To capture the importance of the BSE food scare, we use an index of media coverage based upon a count of newspaper articles per month

⁵ Since we use the generalised impulse response function (Koop *et al.* 1996, Pesaran and Shin, 1998) results are also invariant to the ordering of the variables in \mathbf{x}_{1t} (see Garratt *et al.*, 2003).

⁶ Details regarding the collection and transformation of data are in MAFF (1999).

on the food and health related issues in four national quality newspapers.⁷ Whilst not exclusively about the crisis, BSE and its implications for the safety of beef consumption dominates the index, although reports about other scares such as E Coli 157 and related issues such as abattoir hygiene and cholesterol are also recorded. Supply side shocks of BSE are incorporated in a variable called Net Withdrawals, (NW_t) which represents the sum of net exports (of live cattle, fresh and frozen beef) and cattle removed from the food chain as part of the UK Government's official cull of old and infected cattle. The data are expressed in thousand tonnes of carcass weight equivalent.

5. Empirical Results

As an initial step, the data are tested for the order of integration. The series used comprise 132 monthly observations on retail prices of beef, pork, lamb (RB_t , RP_t and RL_t respectively), the producer price of beef (PB_t), the Meat Scares Index (s_t) and net withdrawals of beef from the UK market (NW_t).⁸ The ADF results are reported in Table 1 and confirm that the data series are non-stationary in levels and stationary in first differences.⁹

Using these data, equation (10) is estimated for $p = 1, \dots, 5$ unrestricted seasonals and intercepts restricted to the cointegration space. The Akaike Information Criterion selects a VAR(2) model, and diagnostic testing for residual auto-correlation, ARCH, and heteroscedasticity does not suggest departure from stated assumptions at the 5% level using either vector or equation-based tests. The null of normally distributed residuals is however strongly rejected owing to the presence of outliers in most of the equations around April 1996, corresponding to the Ministerial announcement in the UK linking BSE and vCJD.

Cointegration Analysis

Cointegration results, reported in Table 2, show two large eigenvalues, pointing to the presence of two equilibrium (stationary) relations among the variables. The formal cointegration tests (Johansen 1988) confirm the presence of two cointegrating relationships at the 5% level, as indeed visual inspection of the cointegrating residuals might suggest (see Appendix)¹⁰.

⁷ The newspapers are *The Times*, *Sunday Times*, *Guardian*, *Observer*. The count is compiled by Euro-PA Associates of Northampton, UK (www.euro-pa.co.uk).

⁸ Being a potentially important substitute for beef, the retail price of chicken was included but was found to be stationary about broken mean at the 1% significance level according to Perron (1989) tests. Another potentially important variable, marketing costs (proxied here by unit labour costs in UK manufacturing), was found to be I(1), but both were found to be redundant in the cointegration analysis and are excluded from the models reported here.

⁹ Application of Perron (1989, 1997) tests to the price series used here, and reported in Sanjuán and Dawson (2003) confirm the presence of a unit root in all price series when allowing for a break in both level and level and trend. Similar testing on the shifters, NW_t and s_t , arrives at the same conclusion.

¹⁰ In the absence of any evidence of a break in the level or trend of these residuals during 1996 we do not apply the Johansen *et al.* (2000) cointegration tests that allow for structural breaks.

Table 1: Augmented Dickey-Fuller test statistics

Variable	Levels (lag)	Differences (lag)	Inference
RB_t	-1.73 (0)	-10.98** (0)	$RB_t \sim I(1)$
RP_t	-1.65 (0)	-9.87** (0)	$RP_t \sim I(1)$
RL_t	-2.23 (3)	-7.39** (2)	$RL_t \sim I(1)$
PB_t	-2.47 (2)	-6.63** (0)	$PB_t \sim I(1)$
NW_t	-3.17 (7)	-6.33** (10)	$NW_t \sim I(1)$
s_t	-2.76 (7)	-5.87** (0)	$s_t \sim I(1)$

Notes: Lag length of the ADF regression is selected according to the Akaike Information Criterion and reported in parentheses adjacent to test statistic; the Augmented Dickey Fuller regression includes a constant and trend (and seasonals for lamb) for the levels and constant (and seasonals for lamb) in differences; critical values derived by MacKinnon (1991); 5% significance denoted by *, 1% by **.

Table 2: Cointegration Test Statistics

Eigenvalues:	0.37	0.28	0.17	0.05	0.03	0.0
H_0	Trace	5% c.v		Maximal Eigenvalue	5% c.v.	
$r = 0$	136.3**	102.1		59.3**	40.3	
$r = 1$	77.0*	76.1		42.0**	34.4	
$r = 2$	35.0	53.1		23.6	28.1	
$r = 3$	11.4	34.9		7.3	22.0	
$r = 4$	4.1	20.0		3.9	15.7	
$r = 5$	0.3	9.2		0.3	9.2	

Notes: Critical values are asymptotic, derived by Osterwald-Lenum (1992); ** and * denote significance at 1% and 5% respectively.

In the absence of additional restrictions, the long-run relations that have been detected are unidentified and merely represent statistical rather than meaningful economic relationships. However, given the discussion in section 3, it is reasonable to assume that they represent the pricing relationships given by equations (6), (7) and (8), of which we would only expect to find two because the margin (8) is the linear combination of (6) and (7). In other words, any two of these equations will by definition determine the third. While the choice is arbitrary as to which two equations to identify, we specify a model in which producer prices are excluded from the first (retail) relation and enter the second (margin) relation with a unit coefficient. This yields the following (t ratios in parentheses):

$$RB_t = 37.48 + 0.88RP_t + 0.67RL_t - 24.09s_t + 0.30NW_t \quad (15)$$

(0.35) (4.89) (1.56) (-4.92) (0.17)

$$(RB_t - PB_t) = -0.46 - 0.32RP_t + 0.53RL_t + 10.29s_t - 0.17NW_t \quad (16)$$

(-0.01)(-1.14) (1.23) (2.06) (-1.50)

Inspection of these suggest that they do indeed represent empirical analogues of the retail and margin relationships as described in (6) and (8). Specifically, note how the price of substitutes (RP_t and RL_t) appear to be relevant in (15), the retail model, but not so in (16), the margin or price transmission relationship. Dropping insignificant regressors yields a final pair of cointegrating vectors given by:

$$RB_t = 0.74RP_t + 0.99RL_t - 30.07s_t \quad (17)$$

(3.22) (4.13) (-3.63)

$$(RB_t - PB_t) = 33.71s_t - 3.86NW_t \quad (18)$$

(6.88) (-3.64)

A likelihood ratio test supports the over-identifying restrictions embodied in (17) and (18) at the 5% level [$\chi^2(5) = 9.1$; p -value = 0.11] indicating that the model represents a congruent simplification of the data. All variables are correctly signed in both equations. Equation (17) describes the retail relationship and shows that as the prices of substitute goods rise, so does the retail price beef, in a manner indicative of substitution. Hence we would expect the presence of substitutes to partially offset the impact of the BSE crisis on price adjustment. The meat scares index (s_t) enters equation (17) as a retail demand ‘shifter’, akin to X in the theoretical model. The empirical results show that consumer concerns over the safety of meat, as measured by media activity, have a negative impact on the retail price of beef. However, because demand shocks are greater at producer rather than retail level, the effect of the index on the margin is to widen it, as can be seen in (18).

Net withdrawals, NW_t enters the margin equation (18) like N in the theoretical model, and as a supply ‘shifter’, it reduces the margin by raising producer prices by more than retail prices. In other words, a decrease in net withdrawals (such as an export ban) will lead, *ceteris paribus*, to an increase in supply on the domestic market and hence prices will fall.

Note that in the price margin equation (18), the two exogenous variables representing the shock to the farm supply function and the retail demand function respectively, are statistically significant at conventional levels.¹¹ Given our discussion following the derivation of equation (8), this implies that market power (either in the form of oligopoly or oligospony power or both) characterises the UK food sector. As such, the results suggest that we cannot reject the hypothesis that market power played a role in influencing the impact of the BSE crisis on the retail-farm margins in the latter half of the 1990s.

The extent to which demand shocks at the retail level are passed back to farmers can be derived by re-writing (18) in terms of producer prices which gives:

$$PB_t = RB_t - 33.71s_t + 3.86NW_t \quad (19)$$

¹¹ Note that t ratios are only asymptotically valid under cointegration although likelihood ratio exclusion tests confirm the statistical significance of these variables to at least one of these cointegrating relation in the system at the 1% level.

The negative coefficient on s_t shows that media activity also depresses producer prices *ceteris paribus*. Also, changes in NW_t that lead to a reduction in domestic supply via exports or by culling raise the producer price of beef.

Impulse response analysis

Given that we have detected the presence of market power through the significance of the shifters, the next step is to examine the specific effects of each shock on beef prices. Prior to reporting the results from this exercise it is worthwhile to note that up to this point it has been assumed that all prices and shifters are potentially endogenous. If the shifters are actually exogenous for the estimation of the long run parameters, estimation efficiency may be improved if they are treated accordingly (Pesaran *et al*, 2000). Whilst it is clear that shocks to the shifters are logically prior to the prices and hence contemporaneously causal for them, the issue of long-run weak exogeneity is largely an empirical issue. It is conceivable that the evolution of prices feeds back on to the shifters, affecting trade (the supply shifter) or the topicality of media interest in health and safety issues (the demand shifter). Using the restricted model given by (17) and (18) the null of weak exogeneity is firmly rejected for each shifter confirming the endogeneity of the shifters in the system.¹²

The generalised impulse response function developed by Koop *et al.* (*op. cit.*) and Pesaran and Shin (*op. cit.*), which explicitly allows for the dynamic interactions between the variables in a system following a specific shock, offers a convenient tool with which to investigate what might be more appropriately called ‘long-run’ responses – the eventual impact that one might observe following a shock to one of the variables. Given the discussion in section 4, we treat the shocks to the shifters as contemporaneously causal for meat prices and Figure 3 shows the simulated effect of a shock of typical size (one standard error, or 79% of the mean) to the meat scares index on all meat prices in the twelve months following this hypothetical shock¹³.

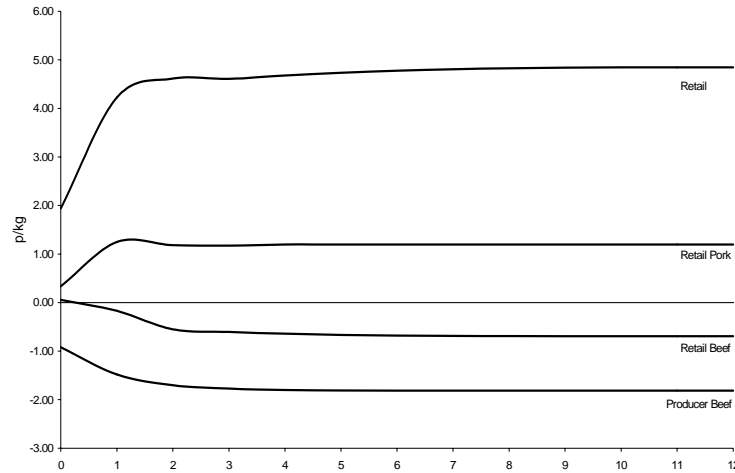
There are two obvious outcomes from the impact on prices following the food scare. First, shocks to the meat scares index leads to a decline in the price of beef whereas the prices of substitute meats rise. Estimates suggest that the retail prices of pork and lamb rise by 0.8% (1.20 p/kg) and 1.8% (4.85 p/kg) respectively, while retail beef prices fall by around 0.3% (0.70p/kg) following a one standard error shock. This would seem to imply that following the BSE scare, consumers reduced their demand for beef, but increase demand for substitute meats. In effect, the results suggest that consumers do not simply stop buying beef following heightened concerns about its safety, but rather they switch at least part of their beef consumption into lamb and pork. Whilst these ‘knock-on’ effects are unsurprising, their quantification underscores the usefulness of impulse response analysis in the inter-related market setting, since they cannot be inferred directly from estimates from the cointegrating

¹² For each shifter (i), the null of weak exogeneity is tested using a likelihood ratio statistic proposed by Johansen (1992) of the restrictions $\alpha_{i,1} = \alpha_{i,2} = 0$. The resulting $\chi^2(2)$ test statistics (of 26.9 and 22.0) reject the null at the 1% level confirming their endogeneity to the long-run parameters in β .

¹³ Individual impulse response figures including the associated confidence intervals are available upon request.

regressions.¹⁴ The second result to note is the differential effect on beef prices at the retail and farm stages. Thus while the retail price of beef falls by around 0.7p/kg, the farm gate price of beef falls by 1.81p/kg, suggesting a ‘pass-back’ coefficient of 2.59. Clearly, shocks at the retail level have far greater impact on farmers than retailers.

Figure 3: The Simulated Dynamic Effect of a (one standard error) Shocks to the Meat Scares Index



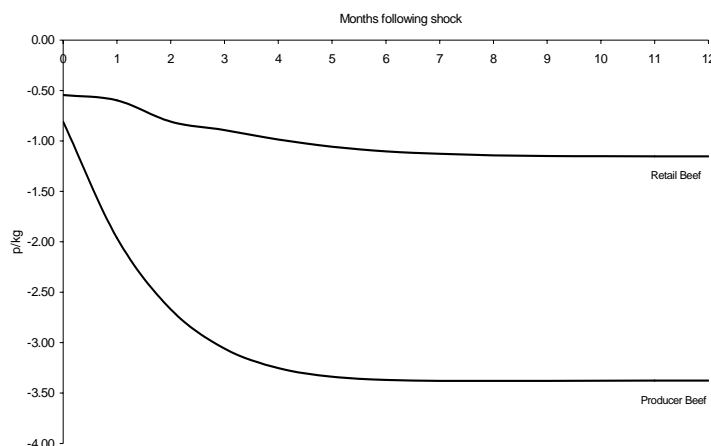
On the supply side, BSE impacted on the national herd via the international ban on UK exports of beef cattle and products and the cull of infected and older cattle. Figure 4 charts the simulated effect of a typical (i.e. a one standard error, $-6,458$ tonne) shock on the retail and producer price of beef. Both prices decline following the reduction in exports but farm gate prices fall 3.38p/kg whereas retail prices fall by 1.15p/kg, implying a ‘pass-through’ elasticity of 0.34. These results are consistent with our observation of (18) that market power will affect the retail-farm spread. If market power did not matter, the demand shock would not have a differential effect on farm and retail prices (see equation 8)¹⁵.

The empirical results relating to the UK beef sector and transmission between retail and farm prices highlight three important issues in mapping out the appropriate econometric strategy when market power may be playing a role. The first is to draw a link between theory and the econometric strategy. Specifically, if market power is likely to be important, this will determine the nature of the cointegrating relationship. Second, if cointegration in the presence of shifters is identified, one can formally detail the structure of the cointegrating equation. Third, only then is it appropriate to investigate the impulse response functions mapping out the dynamics between retail and farm level prices. The results for the beef sector outlined above highlight the relevance of applying this methodology.

¹⁴ Given the strength of the substitution between the various red meats, the partial food scare elasticity of retail beef prices (0.118 at mean values) gives a poor estimate of the total elasticity (of 0.003 at mean values).

¹⁵ To put these price transmission elasticities in context, with perfect competition the ‘pass-through’ elasticity should equal the share of the raw commodity in the industry cost function while the ‘pass-back’ elasticity should reflect the reciprocal of the share of the raw commodity in the industry cost function. Under fairly reasonable assumptions, imperfect competition serves to reduce the pass-through elasticity while increasing the pass-back elasticity.

Figure 4: The Simulated Dynamic Effect of a (one standard error) Fall in Net Withdrawals



As noted in the introduction, we have collected data for 9 other sectors in the UK. Results for other commodity sectors are consistent with the detailed results we have shown for the beef sector and highlight the significance of our approach¹⁶. As such, the procedure for the beef sector outlined above can be applied to these sectors with the resulting impulse response functions having a firm grounding in the theoretical models that highlight the role of market power in the food sector.

6. Summary and Conclusions

This paper has focused on tying empirical studies of price transmission with the underlying theory that highlights a role for market power in determining the outcome. The departure point has related to econometric studies of price transmission that focus on price pairs only, our premise being that cointegration tests relating to price pairs only are inappropriate for addressing whether market power is likely to play a role. We have shown that there is a tie between theory and the appropriate specification of the cointegrating relationship which provides a formal basis for addressing the role of market power. It is only with the appropriate specification that analysis of imperfect price transmission via impulse response functions becomes meaningful and that directly ties theory with the empirics. We have applied this framework to 10 commodity sectors in the UK, most of which confirm the importance of our approach. We have detailed the relevance of our approach with a detailed analysis of price transmission in the UK beef sector. The results suggest that market power in the UK food sector is likely to have played a role in determining price transmission between retail and farm prices. This is consistent with the overall concerns of the Competition Commission but is inconsistent with the recent report commissioned by DEFRA covering the same subject. However, given the lack of formal structure underlying the framework applied in the latter, we doubt whether the results reported there can be used to draw the conclusions that have been reported. In contrast, our approach is both tied to theory and hence is more meaningful in interpreting the overall results. In addition, it is straightforward to apply.

¹⁶ A summary of these results will be provided in an extended paper and commented on in the presentation.

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