The Impact of the Food Scares on Price Transmission in Inter-Related Markets

Tim Lloyd[a]
Steve McCorriston[b]
Wyn Morgan[a]
Tony Rayner[a]

[a] School of Economics, University of Nottingham, Nottingham, NG7 2RD, England, UK
[b] School of Business and Economics, University of Exeter, Exeter, EX4 4PU, England, UK

Contact: tim.lloyd@nottingham.ac.uk

Draft of August 2003

Paper prepared for presentation at the 25th International Conference of Agricultural Economists, August 16-22, 2003, Durban, South Africa

The Impact of the BSE Crisis on Retail-Farm Spreads in the UK

Abstract

This paper is concerned with the impact of the BSE crisis in the UK and focuses on the spread between retail and farm prices. From a theoretical perspective we show that if market power has an effect on the spread between retail and farm prices then this determines the specification of the co-integrating relationship. We also account for different aspects of the BSE crisis including the shift in the retail demand function as well as the shift in the farm supply function due to the mutually opposing effects of a worldwide export ban and the cull of infected and older beef cattle. The empirical results suggest that the impact of the BSE crisis on farm prices to be more than double that of retail prices, the main cause of the change in the retail-farm price spread being the shift in the consumer demand function. Our calculations suggest that the numbers removed via the cull offset the loss of exports, so that the net effect of the BSE crisis on prices is due to the depressed stated of demand. In addition, the results also suggest that we cannot reject the importance of market power in the UK food sector in influencing the retail-farm spread following the outbreak of BSE. This result corroborates the findings of the UK's competition authority, whose recent report found supermarkets undertaking certain practices that were judged to be against the public interest.

Keywords: BSE crisis; price adjustment; impulse response functions

Introduction

In recent years, there has been growing concern about the health and safety of food. In some extreme cases, this has resulted in a food scare, with perhaps the most well-known food scare in recent years being the BSE crisis in the UK. Whilst known since 1986, the BSE crisis erupted in the UK in March 1996 following a Ministerial announcement to Parliament suggesting a link between bovine spongiform encephalopathy (BSE) and the invariably fatal human disease, variant Creutzfeld-Jakob disease (vCJD). To date, 129 deaths can be linked to vCJD in the UK. The announcement led to an immediate 40% fall in the consumption of beef in the UK, and the complete loss of all export markets (including all Member States of the European Union) worth an estimated $1.7bn in 1995 (DTZ Pieda Consulting). In response to the crisis the UK government introduced a large number of measures to lessen the losses to the beef industry and avoid BSE entering the downstream food sector, the most significant of which was the national cull of infected and older cattle, which to date has removed 4.9million cattle from the food chain.

Apart from the obvious concern with human health, an additional concern has been raised that the BSE crisis had a differential effect on UK retailers and farmers. Specifically, it was argued that while the BSE crisis reduced the consumption and price of beef, the decline in beef prices at the retail level was substantially less than that faced at the producer level resulting in a substantial increase in the retail:producer price spread. In this regard, attention was drawn to market concentration at the retail level, the 5-firm concentration ratio in UK food retailing being around 67 per cent (Dobson Consulting; Cooper). The issue was one of the primary reasons for the investigation by the UK anti-trust authorities, the Competition Commission, which 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)

Whilst there have been a number of studies and influential reports in to BSE, most notably the 16 volume report of the Royal Commission into BSE itself (BSE Inquiry), there has been little formal analysis of the potential differential impact of BSE on retail and farm-gate prices, the exception being Lloyd *et al.* who reported the result that beef prices at various stages of the food chain responded differently as a result of the BSE crisis. This paper extends that analysis in three important ways. First, we formally highlight the links between market power and price transmission following shocks to the retail demand and farm supply functions. We show that if market power (in the form of oligopoly and/or oligopsony power) exists then this will influence the specification of the co-integrating relationship describing price transmission, which we estimate. Second, as noted above, the BSE crisis was multi-faceted in that while there was a substantial shift in the retail demand function as consumers reduced their consumption of beef, there were also significant changes in the supply side due to the cull of the 30 month or older cattle and the export ban, both of which began in March 1996 and remained in force throughout the period analysed in this study. We account for these aspects of the BSE crisis in our econometric model and gauge their relative importance on the retail-farm price spread. Third, in determining the impact of BSE on retail and producer markets, we also acknowledge the potential links between beef and substitute meats at the retail level. This is an important issue not least since both prices and consumption of substitutes rose in the wake of the BSE crisis. Clearly, in fully evaluating the impact of a food scare on a particular marketing chain, it is important to account both for the vertical links between stages in the marketing chain of direct concern as well as the horizontal links between related markets at the retail level, since these could also have an impact on the overall assessment. In this regard, the results show that the BSE crisis caused prices at the producer level to fall by almost three times that of retail prices, in a manner consistent with the concerns raised by the UK anti-trust authorities that price transmission would have been influenced by market power in the UK food sector.

The paper is organised as follows. In section 1 we summarise briefly the literature on food scares and how the focus of this paper differs from this extant literature. We also present an overview of the development of the UK BSE crisis over the 1990s. The theoretical framework for identifying the varying impact on retail and farm-gate prices is developed in section 3. This forms the basis for the econometric strategy that is outlined in section 4 while in section 5 we present the empirical results. In section 6 we summarise and conclude.

1. Food Scares: Related Literature and UK Experience

Related Literature

The literature dealing with food scare events is relatively thin, though early research can be dated back to the 1960s (see for example Brown)[1]. More recent is the paper by Smith *et al*. who analyse the impact of contaminated milk in Oahu, Hawaii in 1982. In

---

[1] There is of course a broader and related literature relating to food safety issues. Here our focus is on a well-defined and highly publicised food scare with a specific and serious outcome rather than a food safety issue that may lead to a healthier lifestyle. Examples of related papers on food safety include Brown and Schrader on cholesterol and diet due to the consumption of eggs, Chang and Kinnucan who extend the study of the cholesterol issue to the consumption of butter and Kinnucan *et al*. on the health

their study, they focus upon the loss of sales due to contaminated milk and derive the appropriate level of compensation to producers. Focusing on the UK experience, Burton and Young investigate the impact of the BSE crisis on meat demand.

Also of relevance to the present paper is the way in which the food scare is measured and the methodology used. Smith *et al*. compile an index of press stories relating to the food scare in question. This index (separated into positive and negative news stories) is then incorporated into a regression model to evaluate the impact of the contamination event on sales. With the focus also on quantities (this time in terms of consumption), Burton and Young use an Almost Ideal Demand system to evaluate the impact of BSE where the impact of BSE is proxied by the use of a media index. Lloyd *et al*. also consider the impact of BSE in the UK meat sector, although their attention focuses upon price adjustment rather than the specification of a meat demand function[2].

BSE in the UK

Detailed accounts of the UK's experience of BSE are available elsewhere (*see inter alia,* BSE Inquiry; DTZ Pieda Consulting) and thus we merely sketch some of the key aspects of the crisis here. The existence of BSE in cattle was first identified in the UK in the 1980s, although at that time, it was not thought to have implications for human health. This view was challenged in the 1990s when it was discovered that that BSE could jump species following the death of a cat from BSE symptoms. However, there was continued re-assurance from the UK government and its Chief Medical Officer

issues and meat consumption. Ippolito and Mathios focus on advertising of health attributes and the consumption of ready-to-eat cereals.

[2] Other studies of the BSE crisis include MacDonald and Roberts who undertake a computable general equilibrium study and Henson and Mazzocchi who apply an event study to firms in the agribusiness.

that British beef was safe to eat. When the first human death from vCJD occurred in 1995, official confirmation of the possible link between vCJD and BSE was announced. To date there have been 129 deaths due to vCJD though the expectation is that this number could rise dramatically in the future particularly as vCJD has a long incubation period (New Scientist).

Following the discovery of the link between BSE and vCJD, consumption of beef fell immediately by 40 per cent. In addition, there were significant policy interventions that also affected the market. First, a European Union-wide ban on all UK beef sales was imposed in March 1996, in effect meaning British farmers faced an export ban. Net exports of beef, which stood around 39,000 calves and 23,000 tonnes of fresh and frozen carcass beef per month, ceased at this point, although it is interesting to note that imports did not change markedly after the ban. This reluctance to consume beef – even from BSE free sources- reflects the commonplace 'over-reaction' to food-based risks (Kinnucan et. al)[3]

Second, and contemporaneous with the export ban, was the imposition of a culling order which consisted of three elements: an "Over 30 Month" cull where all cattle over this age were culled; a "Selective Cull" where known infected cattle were culled; and an "Offspring Cull" where the young of infected cows were culled. The "Over 30 Month Scheme" represented approximately 98% of the 4,900,000 cattle destroyed during the sample period. The ban was lifted in 2003 although the cull is ongoing. Both were in force during our sample period.

[3] Although unknown, the probability of contracting vCJD from BSE-infected meat was considered to be minuscule, yet the media interest at the time offered a very different impression to consumers.

2. The Impact of Food Scares on Price Transmission

As noted in the introduction, one of the main concerns in the UK in the latter half of the 1990s was that the BSE crisis appeared to have a more significant impact on farmers than on food retailers. As a consequence, one of the main aspects of the Competition Commission enquiry into the food retailers was that market power at the retail level resulted in smaller declines in retail beef prices compared to those faced by farmers. In this section we outline a framework that is consistent with this argument. In addition, we also pay some attention to the role of related markets for substitute meats in determining the impact on the price transmission effect of the food scare.

There is a broad literature on the issue of price transmission between the retail and farm levels and what factors may influence it[4]. The most notable early paper on this issue was by Gardner 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, McCorriston *et al*. (1998) showed 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 McCorriston *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 spread and hence price transmission are likely to be oligopsony power (Lloyd *et al*.), and the source of the exogenous shock (i.e. whether the shift occurs in the retail demand or farm supply function (Gardner). Related to this, the nature of the

[4] The following discussion is not aimed at a comprehensive review of the price transmission issue and determinants of retail-farm spreads. For a more detailed discussion, see Wohlgenant.

shift in the demand or supply function is likely to take (i.e. whether it is an intercept or rotation of the relevant curve) will also have an impact[5].

In the theoretical framework outlined below, we draw on this literature to highlight the potential role of market power on the retail-farm spread. However, rather than derive an explicit price transmission elasticity, we use the framework to determine our econometric strategy. Explicitly, in the spirit of Holloway and Reed and Clark, 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 spread equation. If market power is not a feature of the sector, then there would be no *a priori* case for their inclusion. Therefore, while we do not retrieve an explicit measure of market power, in terms of deriving the relative effect of BSE on retail and farm prices, market power will influence the outcome if the demand and supply shifters are found to be statistically significant in the reduced form model.

Economic Framework

The demand function for the processed product is given by:

$$Q = h(R, R^s, X) \tag{1}$$

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):

$$P = k(A, N) \tag{2}$$

[5] These studies focus primarily on price transmission. There are a number of studies that also relate to this issue such as the impact of oligopoly and oligospony power on producer welfare (e.g. Sexton) and the form of the supply curve shift (Alston et al.).

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:

$$\pi_i = R(Q)Q_i - P(A)A_i - C_i(Q_i) \tag{3}$$

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 follows[6]. The first-order condition for profit maximisation is given by:

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

In order to get an explicit solution, consider linear functional forms for equations (1) and (2) and assume $a = 1$ (which is consistent with the construction of the data series):

$$Q = h - bR + eR^s + cX \tag{1'}$$

$$P = k + gS \tag{2'}$$

with domestic supply being given by:

$$S = Q + N$$

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

[6] Note that we are not pre-judging the form of technology that links these two sectors. However, the nature of the technology was not a specific issue in the Competition Commission investigation into the food retailers. It should also be observed that with a Leontief technology, the nature of the exogenous shock is important in identifying market power. However, we are not interested in identifying market power parameters explicitly but only if the presence of market power would have been likely to influence the retail-farm spread over the BSE crisis period.

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

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[7]. Using (1'), (2') and (4'), we can derive an explicit solution for the endogenous variables:

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

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

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

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

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

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:

$$R - P = M \qquad \textbf{(9)}$$

i.e. the source of the retail-farm price spread in a perfectly competitive industry is due to marketing costs. 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 chain is important, then they will influence the spread between retail and farm prices i.e. they will influence these prices by differential amounts.

Note also that the above framework includes the role of substitutes at the retail level in affecting the outcome of the BSE crisis. In general form, the impact of a food scare on retail prices can be given by:

$$\frac{\partial R}{\partial X} = \frac{\partial R}{\partial X} + \frac{\partial R}{\partial R^s}\frac{\partial R^s}{\partial X}$$

The first argument on the right-hand side is negative. A food scare shifts the demand curve to the left (i.e. an increase in X is equivalent to negative impact on demand) and hence reduces retail prices. The second argument is potentially ambiguous and will depend on how consumers respond to meat consumption as a whole. If $\partial R^s / \partial X$ is positive (negative), then the presence of substitute meats will offset (exacerbate) the impact of the food scare on retail prices.

Equations (6) and (8) form the basis of our econometric strategy. Using equation (6), we derive the direct effects of the demand and supply shifters on retail prices. Using

[7] As noted above, the retail food sector in the UK is highly concentrated.

equation (8), we focus on 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 not enter our econometric model. Hence 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 gate prices to varying degrees. Consequently, we can use (6) and (8) to derive the impact of the BSE crisis on retail and farm prices with N capturing the farm-level shifts associated with the export ban and the cull of cattle and X capturing the impact of the food scare at the retail level.

Econometric Methodology

General

Applied to the current context, the theoretical model set out above demonstrates the differential effect on producer and retailer beef prices following the BSE crisis in the UK. Since prices are likely to be non-stationary and co-integrated, it is appropriate to couch the empirical analysis in a vector autoregressive (VAR) framework (Hendry and Doornik, p129). Consider a VAR(p) model:

$$\mathbf{x}_t = \mathbf{\Phi}_1 \mathbf{x}_{t-1} + \mathbf{\Phi}_2 \mathbf{x}_{t-2} + \ldots + \mathbf{\Phi}_p \mathbf{x}_{t-p} + \mathbf{\Psi} \mathbf{w}_t + \boldsymbol{\varepsilon}_t \qquad (10)$$

where $\mathbf{x}_t$ is a ($m \times 1$) vector ($1,2,\ldots,i, j, \ldots,m$) of jointly determined I(1) variables, $\mathbf{w}_t$ is a ($q \times 1$) vector of deterministic and or exogenous variables and each $\mathbf{\Phi}_i$ ($i = 1,\ldots, p$) and $\mathbf{\Psi}$ are ($m \times m$) and ($m \times q$) matrices of coefficients to be estimated using a (t = 1, . . .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, $\mathbf{\Sigma}$.

The error correction representation of (10) is observationally equivalent but facilitates estimation and hypothesis testing since all terms are stationary (Hendry and Doornik, p60). This re-parameterisation is given by:

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

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 $\boldsymbol{\Gamma}_i$ coefficients in (17) estimate the short-run effect of shocks on $\Delta\mathbf{x}_t$, and thereby allow the short and long-run responses to differ.

In the current context, $\boldsymbol{\beta}$ represents the linkages that bind the prices together in the long run. As section 2 demonstrates, these linkages occur either across substitutes at the retail level or between marketing stages for a single good. However, as Lütkephol and Riemers make clear, despite offering estimates of the 'long run' they are by construction partial derivatives predicated on the *ceteris paribus* assumption. When the variables in a co-integrated 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. Interpretation can be confounded further when more than one co-integrating vector is present in the system (Lütkephol).[8] If what is actually wanted is an estimate of what might happen following a specific shock, then impulse response analysis

[8] An illustration of this problem is given in the empirical analysis which contains two co-integrating vectors.

which takes account of these interactions provides a tractable and potentially attractive solution.

Expressing the VAR in its moving average [MA(∞)] representation, we have:

$$\mathbf{x}_t = \boldsymbol{\varepsilon}_t + \mathbf{A}_1\boldsymbol{\varepsilon}_{t-1} + \mathbf{A}_2\boldsymbol{\varepsilon}_{t-2} + \ldots + \sum_{i=0}^{\infty}\mathbf{A}_i\boldsymbol{\Psi}\mathbf{w}_{t-i} \qquad (12)$$

where the ($m \times m$) coefficient matrices $\mathbf{A}_i$ are obtained according to:

$$\mathbf{A}_i = \boldsymbol{\Phi}_1\mathbf{A}_{i-1} + \boldsymbol{\Phi}_2\mathbf{A}_{i-2} + \ldots + \boldsymbol{\Phi}_p A_{i-p} \qquad i = 1,2,\ldots,$$

with $\mathbf{A}_0 = \mathbf{I}_m$ and $\mathbf{A}_i = \mathbf{0}$ for $i < 0$. In this moving average representation of the model:

$$\mathbf{A}_n = \frac{\partial \mathbf{x}_{t+n}}{\partial \boldsymbol{\varepsilon}_t'}$$

comprises coefficients that measure the effects *n* periods after a system-wide shock to the disturbances on each variable in the system. Specifically, the ij^{th} element of the $\mathbf{A}_n$ matrix defines the effect of a one unit increase in the j^{th} variable's disturbance at time *t* (ε_{jt}) on the i^{th} variable at time $t+n$, ($x_{i,t+n}$), holding all other disturbances at all dates constant. A plot of the ij^{th} element of $\mathbf{A}_n$, i.e.,

$$\frac{\partial x_{i,t+n}}{\partial \varepsilon_{jt}} \qquad (13)$$

as a function of n is the *impulse response function* $\{\varphi_{i,j}(n)\}$. However, the *ceteris paribus* clause upon which (13) is predicated, denies the dynamic interaction that the VAR attempts to capture. Indeed, (13) assumes that the disturbances in (10) are un-correlated and thus that $\boldsymbol{\Sigma}$ is diagonal. Given that a shock to ε_{jt} may simultaneously perturb all other variables, and in turn $\mathbf{x}_{t+n}$, the total effect of the shock is the quantity of interest. Focussing on $x_{i,t+n}$, the net effect of these interactions is given by:

$$\frac{dx_{i,t+n}}{d\varepsilon_{jt}} = \frac{\partial x_{i,t+n}}{\partial \varepsilon_{1t}}\delta_1 + \frac{\partial x_{i,t+n}}{\partial \varepsilon_{2t}}\delta_2 + \ldots + \frac{\partial x_{i,t+n}}{\partial \varepsilon_{mt}}\delta_m \tag{14}$$

where $(\delta_1, \delta_2, \ldots, \delta_m) = \boldsymbol{\delta}$ defines the extent to which the shock to ε_{jt} contemporaneously 'impacts' on all other variables. The time profiles describing the effect of a hypothetical shock on the level of $\mathbf{x}_t$, take into account the 'knock-on' as well as the 'feedback' effects that characterise the inter-relationships between variables in a dynamical system such as (10). The *generalised impulse response function* (Pesaran and Shin) offers a measure of (14) that, unlike the more usual orthogonalised impulse response function of Sims, is invariant to the ordering of the variables in the VAR and is thus unique.

In what follows we use the generalised impulse response function 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 and (b) gauge the relative importance of the demand and supply shocks that characterise the BSE crisis[9].

[9] One additional consideration in the econometric strategy is threshold co-integration whereby prices may be characterised by non-adjustment within a pre-determined price band (see, for example, Goodwin and Holt and Miller and Hayenga). These models are useful for identifying the role of transaction costs where, for example, firms may not adjust retail prices for relatively small changes in farm prices due to the costs of doing so. However, such an approach is acceptable for relatively small shocks to the food sector. In the current context, the shocks to the UK food sector due to the BSE crisis were substantive. For example, a typical standard error of the data proxying for the food scare at the retail level was large. Such large variation also characterised the shift in the farm supply function. In addition, most threshold models are applied to bi-variate models with one co-integrating vector. With a six variable model and two co-integrating vectors, and where each variable could have its own threshold parameter, it would be unlikely that the application of the threshold model would give

3. Data

Monthly price data spanning January 1990 to December 2000 are supplied by the UK's Department of Environment, Food and Rural Affairs (DEFRA).[10] Retail price series 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 on the food and health related issues in four national quality newspapers.[11] 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. The

---

meaningful insights. Given these considerations, we did not pursue the threshold approach as it would likely add very little to the analysis presented here.

[10] Details regarding the collection and transformation of data are in MAFF (1999).

monthly count of articles is shown in Figure 1 and the series enters the empirical analysis in log form (s_t).

Figure 1: Article Count of the Health and Safety of Beef (about here)

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[12]. The data are expressed in thousand tonnes of carcass weight equivalent.[13]

Figure 2: Net withdrawals of Beef from the UK Market (about here)

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).[14] The results are reported in

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

[12] Ideally, we would have incorporated these shocks separately. However, the data for each shock is highly collinear which affected the econometric results. The Net Withdrawal variable is therefore a composite of two sources of shock to the farm supply schedule.

[13] The authors are grateful to Tony Fowler of the Meat and Livestock Commission and Ken Addison of the Rural Payments Agency for help and advice in construction of the data series, further details of which are available from the authors upon request.

[14] 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. It also had no statistically significant impact in the cointegration analysis and is excluded from the models reported here.

Table 1 and confirm that the data series are non-stationary in levels and stationary in first differences, as visual inspection of the data (see Appendix) suggests.

Table 1: Augmented Dickey-Fuller test statistics

Variable	Levels (lag)	Differences (lag)	Inference
RB_t	-1.73 (0)	-10.98** (0)	RB_t ~I(1)
RP_t	-1.65 (0)	-9.87** (0)	RP_t ~I(1)
RL_t	-2.23 (3)	-7.39** (2)	RL_t ~I(1)
PB_t	-2.47 (2)	-6.63** (0)	PB_t ~I(1)
NW_t	-3.17 (7)	-6.33** (10)	NE_t ~I(1)
s_t	-2.76 (7)	-5.87** (0)	s_t ~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; 5% significance denoted by *, 1% by **.

Using these data, equation (10) is estimated for $p = 1, \ldots, 5$ unrestricted seasonals and intercepts restricted to the co-integration 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[15]. Plots of actual and fitted values from each of the equations and their corresponding scaled residuals are given in the appendix.

[15] One further issue refers to the changes in agricultural policy that may have affected beef prices. To the extent that policy support policies changed, this would have been affected in the price series we observed. For substantive changes in policy, we therefore may have expected see a structural change in

Co-integration Analysis

Co-integration results, reported in Table 2, show two (or possibly three) large eigenvalues which point to the presence of at least two equilibrium relations among the variables. Both formal tests indicate the presence of two such relationships at the 5% level, as indeed visual inspection of the co-integrating residuals might suggest (see appendix).

Table 2: Co-integration 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; ** and * denote significance at 1% and 5% respectively.

In the absence of additional restrictions, the long-run relations are unidentified and merely represent statistical rather than meaningful economic relationships. However, given the discussion in the preceding sections, it is possible that they may represent the (horizontal) relationship between meat prices at the retail level (see equation 6) and the (vertical) price transmission relationship between retail and producer beef prices (see equation 8). As a first step, an exactly identified model is estimated in which producer prices are excluded from the first (retail) relation and enter the second

the data. However, we could find no evidence of such structural change in our pre-testing of the data and so is not directly incorporated in the model estimated.

(vertical) relation with a unit coefficient. This yields the following (t ratios in parentheses):

$$\underset{(0.35)}{RB_t = 37.48} + \underset{(4.89)}{0.88RP_t} + \underset{(1.56)}{0.67RL_t} - \underset{(-4.92)}{24.09s_t} + \underset{(0.17)}{0.30NW_t} \tag{15}$$

$$(RB_t - PB_t) = \underset{(-0.01)}{-0.46} - \underset{(-1.14)}{0.32RP_t} + \underset{(1.23)}{0.53RL_t} + \underset{(2.06)}{10.29s_t} - \underset{(-1.50)}{0.17NW_t} \tag{16}$$

Note how the price of substitutes (RP_t and RL_t) appear to be relevant in the retail model (15) but not so in the price transmission relationship (16). Incorporating theory relevant restrictions and dropping insignificant regressors yields a final pair of cointegrating vectors given by:

$$RB_t = \underset{(3.22)}{0.74RP_t} + \underset{(4.13)}{0.99RL_t} - \underset{(-3.63)}{30.07s_t} \tag{17}$$

$$(RB_t - PB_t) = \underset{(6.88)}{33.71s_t} - \underset{(-3.64)}{3.86NW_t} \tag{18}$$

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[16]. As such, these equations may be interpreted as the equilibrium relations posited in section 2 characterising the vertical links between retail and farm prices and the role of inter-related markets at the retail level. Equation (17) describes the 'horizontal' retail relationship and shows that as the price of substitute goods rise so does the retail price beef in a manner indicative of substitution. The estimates suggest that, *ceteris paribus*, a one p/kg rise in the retail price of pork (lamb) increases the price of retail beef by 0.74p/kg (0.99p/kg). At mean values, these equate to price response elasticities of 0.41 for pork and 1.04 for lamb.

[16] Note also that the variables have the correct sign. Recall that a positive shift in s_t represents a shift to the left of the retail demand function. The coefficients in the farm-retail spread are also correctly signed which can be seen by re-arranging (18) in terms of *PB* alone, [see equation (19)].

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, and translates to an elasticity of (-30.07/254=) –0.12 at mean values. Thus, following a food scare such as BSE, consumers reduce consumption of beef, leading to a decline in retail prices.

Note that in the price transmission relationship, (18), the two exogenous variables representing the shock to the farm supply function and the retail demand function respectively are statistically significant. 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, we can reject the hypothesis that market power did not play a role in influencing the impact of the BSE crisis on the retail-farm spread in the latter half of the 1990s. This result is in contrast to the studies by Holloway and Reed and Clark who argued that they could not reject the assumption of competitive markets[17]. However, it does corroborate the results of the Competition Commission's inquiry that found a range of selling and purchasing practices commonly used by UK supermarkets to be against the public interest (Competition Commission).

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:

$$PB_t = RB_t - 33.71 s_t + 3.86 NW_t \tag{19}$$

[17] Note also that given the nature of the data used and that the BSE effect was UK-wide, we do not suffer from the aggregation problems that affect the interpretation of the Holloway and Clark and Reed studies.

The negative coefficient on s_t shows that media activity also depresses producer prices *ceteris paribus,* corresponding to an elasticity at mean values of (-33.71/133=) –0.25. This is approximately half as large again as the effect of media activity at the retail level. Also changes in NW_t that lead to a reduction in domestic supply via exports or by culling raise the producer price of beef. The estimate suggests that for every 1000 tonnes withdrawn from the domestic market prices rise by around 4p/kg, implying an elasticity of (3.86*4.61/133=) 0.13.

Whilst these 'long run' estimates are informative, it should be borne in mind that they represent partial derivatives from conditional equations and thus ignore the 'knock-on' and 'feed-back' effects that characterise the inter-relationships between the meat prices in this system. Specifically, it should be recognised that one of the prices in (18) is fixed and thus strictly speaking it is a 'conditional' retail-farm spread equation, i.e. it describes what happen to the mark-up following a change to demand or supply conditional on one of the prices remaining fixed. As such, its value is limited since we wish to determine the effect of these shocks on the price spread explicitly[18]. Thus it is noteworthy that despite being estimates of the 'long run', parameters of the co-integrating regression are generally ill-suited to simulation-type exercises where variables are inter-related (Lütkepohl). In the present case where a change to (say) retail beef demand may be expected to affect prices further up the supply chain (and of substitutes too) in addition to retail price itself, different tools are required.

Impulse response analysis

[18] These feedback effects also arise with respect to the meat scares index as the change in the retail-farm spread during the BSE crisis would also have been reported in the media.

The generalised impulse response function developed by Pesaran and Shin, 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. Figure 3 shows the simulated effect of a shock of typical size (one standard error) to the meat scares index on all meat prices in the twelve months following this hypothetical shock.

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

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 1.20 p/kg and 4.85 p/kg respectively, while beef prices fall by around 0.70p/kg following a one standard error shock. This would seem to imply that following a food (BSE) scare, consumers reduce 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 from estimates from the co-integrating regressions.

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.

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. one standard error –6,458 tonnes) 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).

Figure 4: The Simulated Dynamic Effect of a (one standard error) Shock to Net Withdrawals (about here)

The Impact of BSE on UK Beef Prices

As noted, the BSE episode can be characterised as having both demand- and supply-side effects. The demand side is manifested in the reduction in domestic demand while the supply side is shown through both the ban on exports and the cull of older cattle. Two questions then arise: what is the total impact of these effects on prices in the beef market and what is the relative size of them? We present some (albeit tentative) indication of the answers to these questions[19].

To calculate the supply-side effects we note that monthly exports of fresh and frozen beef which had averaged 22,810 tonnes (cwe) and exports of around 39,200 live calves ceased in April 1996. Equating 40 calves to one tonne of (cwe) beef means that the ban prevented the export of around (22,810+39,200/40=) 23,790 tonnes of (cwe) beef per month. Using estimates from the impulse response analysis suggests that the long-run impact of the ban was to lower average monthly retail prices by 4p/kg and producer prices by 12p/kg.

The impact of the cull requires a conversion from head count figures into tonnes of cwe for consistency. During the 1996(4) to 2000(12) sample period the average number of cattle culled per month was 86,014 and the average weight of cows culled was 556 kg. Thus, the average weight of culled cattle per month was 48,014 tonnes. These figures are live-weight, so need to be adjusted by the 'killing-out percentage' of 0.54. Hence, the monthly cull amounts to 25,928 tonnes of cwe. Using the impulse response estimates again suggest that the long-run impact of the cull was to raise average monthly retail prices by 5p/kg and producer prices by 14p/kg.

Calculation of the demand side effects of BSE are less straightforward, not least because media activity was not in the form of a once-and-for-all shock but was staggered over the entire period. Furthermore, whilst it is possible to accumulate the

[19] Note that whereas the export ban and cull resulted in the abrupt step-change in the amount of beef on the domestic market, the effect on demand was staggered over a long period of time. This means that when we attempt to estimate the impact of BSE on prices, we need to treat the demand and supply side shocks differently.

price effect of these shocks, it seems reasonable to assume that consumers' recent experience of media activity carries more weight than those that occurred in the past. To do otherwise, might exaggerate the effect of media activity on beef prices. In addition, estimates are in terms of proportionate shocks and thus a baseline from which to measure proportionate shocks is required. To allow for these factors we compute a baseline that is a one-year moving average that is updated as new articles are published i.e. we compare the actual count as a proportion of the number of articles appearing in the last 12 months. Mutliplying these shocks by the estimates generated from the impulse response functions (which estimated that a doubling in media activity induces price falls of 0.87p/kg at retail and 2.29p/kg at the producer level) we accumulate the monthly effects over time. Under this scenario, retail prices fall by 19p/kg and producer prices fall by 45p/kg.

The price effects of these factors are summarised in Table 3. Overall, the results show that the impact resulting from the shift in the retail demand function dominated (by a factor of three or four) that of the export ban or cull. Furthermore, for all factors, the impact on producer prices was significantly higher than the impact on retail prices. Indeed, the net effect of the BSE crisis is calculated to be almost three times greater than the impact on retail prices.

Table 3: Estimated Price Impacts of the BSE Crisis (p/kg)

Effect due to:	Retail Prices	Producer Prices
Export Ban	-4	-13
Cull	5	14

Media activity	-20	-46
Net Effect	**-19**	**-45**

Despite the relatively crude nature of the calculations they suggest that the cull completely offset the price effects of the export ban. This arises since the cull removed approximately the same quantity of beef as had been exported prior to the ban. Whilst there is no reason to assume that the number of cattle culled to protect human health should have equalled the volume of lost exports, given that farmers received full market price compensation for the cattle culled, the outcome might be viewed one that met the dual political aim of protecting both beef consumer and producer from the effects of BSE.

Summary and Conclusions

This paper has focussed on the impact of BSE in the UK on prices at both the retail and producer levels. Specifically, it was motivated by the public concerns raised about the differential impact on retailers and producers, the concern being that prices at the farm gate fell by more than retail prices in the wake of the BSE crisis. The principal source of that concern was market power of food retailers which lead to a subsequent investigation by the UK Competition Commission. The contribution of this paper has been two-fold. First, we have shown formally that if market power exists, then exogenous shocks to either the retail or farm supply function will likely have an effect on the retail-farm price spread. Second, since the BSE crisis in the UK can be represented as a combination of demand and supply side shocks, we can gauge the effects of these shocks on retail and farm prices respectively. The economic model we develop forms the basis for our econometric strategy, results from which confirm that shocks to the retail demand and farm supply functions affected the retail-farm spread. While they do not identify the extent of market power, nor whether oligopoly power dominates oligopsony power, the results provide a *prima facie* justification for the issues addressed by the UK Competition Commission i.e. that market power in the UK food sector was a source of concern against the background of the BSE crisis.

Second, the results from the empirical model show that as a consequence of the BSE crisis, producer prices fall by almost three times that of retail prices. Moreover, the effect of the shift in the retail demand function as consumers reduced their consumption of beef and switched into substitute meat products had a more significant impact on retail and producer prices than the shift in the farm supply function due to the effect of the export ban on UK beef or the cull of older cattle. Indeed, it appears that the numbers culled were similar to the number exported prior to the crisis and thus were to all intents and purposes mutually offsetting. As a result the net effect of BSE is solely due to its effect on domestic demand, which is substantial. The overall implication of the analysis presented above is that, while direct concern of food scares obviously relates to the health and well-being of consumers, where market power in the food sector exists, it will also give rise to distributional effects between the retail and farm sectors.

References

Alston, J.M., R.J. Sexton and M. Zhang ‘Imperfect Competition, Functional Forms, and the Size and Distribution of Research Benefits’ *American Journal of Agricultural Economics*, 21 (1999): 155-172.

Brown, J.D. 'Effect of a Health Hazard 'Scare' on Consumer Demand' *American Journal of Agricultural Economics*, 51 (1969): 676-678.

Brown, D. and L. Schrader ' Cholesterol Information and Shell Egg Consumption', *American Journal of Agricultural Economics*, 72 (1990), 548-555.

Burton, M. and Young, T. ‘The Impact of BSE on the Demand for Beef and other Meats in Great Britain’, *Applied Economics*, 28 (1996), 687-693.

Buse, R.C. 'Total Elasticities- A Predictive Device' *Journal of Farm Economics*, 40 (1958): 881-891.

BSE Inquiry. *The Inquiry in to BSE and variant CJD in the United Kingdom*. A Report in 16 volumes. Crown Copyright, London 2000. (http://www.bseinquiry.gov.uk/)

Competition Commission Report on the Supply of Groceries from Multiple Stores in the United Kingdom (3 Volumes). CM 4842. London HMSO.

Cooper, D. ‘Findings from the Competition Commissions’ Inquiry into Supermarkets’ *Journal of Agricultural Economics*, 54(1), (2003), 127-143.

Dobson Consulting. *Buyer Power and its Impact on Competition in the Food Retail Distribution Sector of the European Union* Report prepared for the European Commission. Brussels (1999).

DTZ Pieda Consulting. *The Economic Impact of BSE on the UK Economy*. A Report Commissioned by the UK Agricultural Departments and the HM Treasury. Manchester, March 1998.

Gardner, B. 'The Farm-Retail Spread in a Competitive Industry', *American Journal of Agricultural Economics*, 57 (1975), 399-409.

Gardner, B.L. ‘The Farm-Retail Spread in a Competitive Food Industry’ *American Journal of Agricultural Economics*, 57, (1975): 399-409.

Gardner, B.L. *The Economics of Agricultural Policies*. Macmillan, New York (1987)

Goodwin, B. K. and M. Holt ‘Price Transmission and Asymmetric Price Adjustment in the US Beef Sector’ *American Journal of Agricultural Economics*, 81 (1999): 630-637.

Hendry, D.F. and J.A. Doornik *Modelling Dynamic Systems using PcGive 10* volume II, Timberlake, London (2001).

Henson, S. and M. Mazzocchi 'Impact of Bovine Encephalopathy on Agribusiness in the United Kingdom: Results of an Event Study on Equity Prices' *American Journal of Agricultural Economics*, 84 (2002): 370-386

Holloway, G. J. 'The Farm-Retail Spread in an Imperfectly Competitive Food Sector' *American Journal of Agricultural Economics*, 73 (1991): 979-989.

Ippolito, P.M. and A. D. Mathios 'Information, Advertising and Health: a Study of the Cereal Market' *RAND Journal of Economics*, 21 (1990): 459-480.

Johansen, S. 'Statistical Analysis of Co-integrating Vectors' *Journal of Economic Dynamics and Control*, 12 (1988): 231-254.

Kinnucan, H.W., H. Xiao, C-J Hsia and J.D. Jackson 'Effects of Health Information and Generic Advertising on US Meat Demand', *American Journal of Agricultural Economics*, 79 (1997), 13-23.

Kinnucan, H.W., O. Myrland and L. Paudel 'Relative Impacts of Health Information and Advertising on Commodity Markets: US Meats' in Chern, W.S. and Rickersten K. *Health, Nutrition and Food Demand*, CABI Publishing, Cambridge MA, 2003.

Lloyd, T. A., S. McCorriston, C. W. Morgan and A .J. Rayner 'The Impact of Food Scares on Price Adjustment on the UK Beef Market' *Agricultural Economics* 25 (2001): 347-357

Lloyd, T. A., S. McCorriston, C. W. Morgan and A .J. Rayner 'Market Power and the Impact of Food Scares in the UK' Mimeo, University of Exeter, 2002

Lütkepohl, H. 'Interpretation of Cointegration Relations – Comments on "Estimating Systems of Trending Variables"' *Econometric Reviews*, 13(3), (1994), 391-394.

Lütkepohl, H. and Reimers, H.E. 'Impulse Response Analysis of Co-integrated Systems', *Journal of Economic Dynamics and Control*, 16 (1992): 53-78.

MacDonald, S. and D. Roberts, "The Economy Wide Effects of the BSE Crisis: A CGE Analysis" *Journal of Agricultural Economics*, 49 (1998): 458-471

MAFF (1999) A Report of an Investigation into the Relationship between Producer, Wholesale and Retail Prices of Beef, Lamb and Pork. Ministry of Agriculture, Fisheries and Food, HMSO, London.

MacKinnon, J.G. 'Critical Values for Co-integration Tests' in R.F. Engle and C.W.J.Granger (eds.) *Long-Run Relationships*. Oxford University Press, Oxford (1991)

McCorriston, S. C.W. Morgan, A.J. Rayner ‘Processing Technology, Market Power and Price Transmission’ *Journal of Agricultural Economics*, 49 (1998): 185-201.

McCorriston, S. C.W. Morgan, A.J. Rayner ‘Price Transmission and the Interaction between Market Power and Returns to Scale’ *European Review of Agricultural Economics*, 28 (2001): 143-160.

Miller, D.J. and M. Hayenga ‘Price Cycles and Asymmetric Price Transmission in the US Pork Market’ *American Journal of Agricultural Economics*, 83 (2001): 551-562.

New Scientist 'The Human Tragedy May Just be Beginning', 4 November (2000)

Osterwald-Lenum, M. ‘A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Co-integration Rank Test Statistics’, *Oxford Bulletin of Economics and Statistics*, 56 (1992): 461-472.

Perron (1989) in Rao book

Pesaran, H. and Shin, Y. 'Generalised Impulse Response Analysis in Linear Multivariate Models'. *Economics Letters*, 58 (1998): 17-29.

Reed, A. J. and J.S. Clark ‘Structural Change and Competition in Seven US Food Markets’. US Department of Agriculture, Economic Research Service, Technical Bulletin No. 1881, 2000.

Rogers, R.T and R. J. Sexton 'Assessing the Importance of Oligopsony Power in Agricultural Markets' *American Journal of Agricultural Economics* 76 (1994); 1143-1150.

Sexton, R. J. ‘Industrialization and Consolidation in the US Food Sector: Implications for Competition and Welfare’ *American Journal of Agricultural Economics* 82 (2000): 1087-1104.

Sims, C. ‘Macroeconomics and Reality’ *Econometrica,* 48 (1980): 1-48.

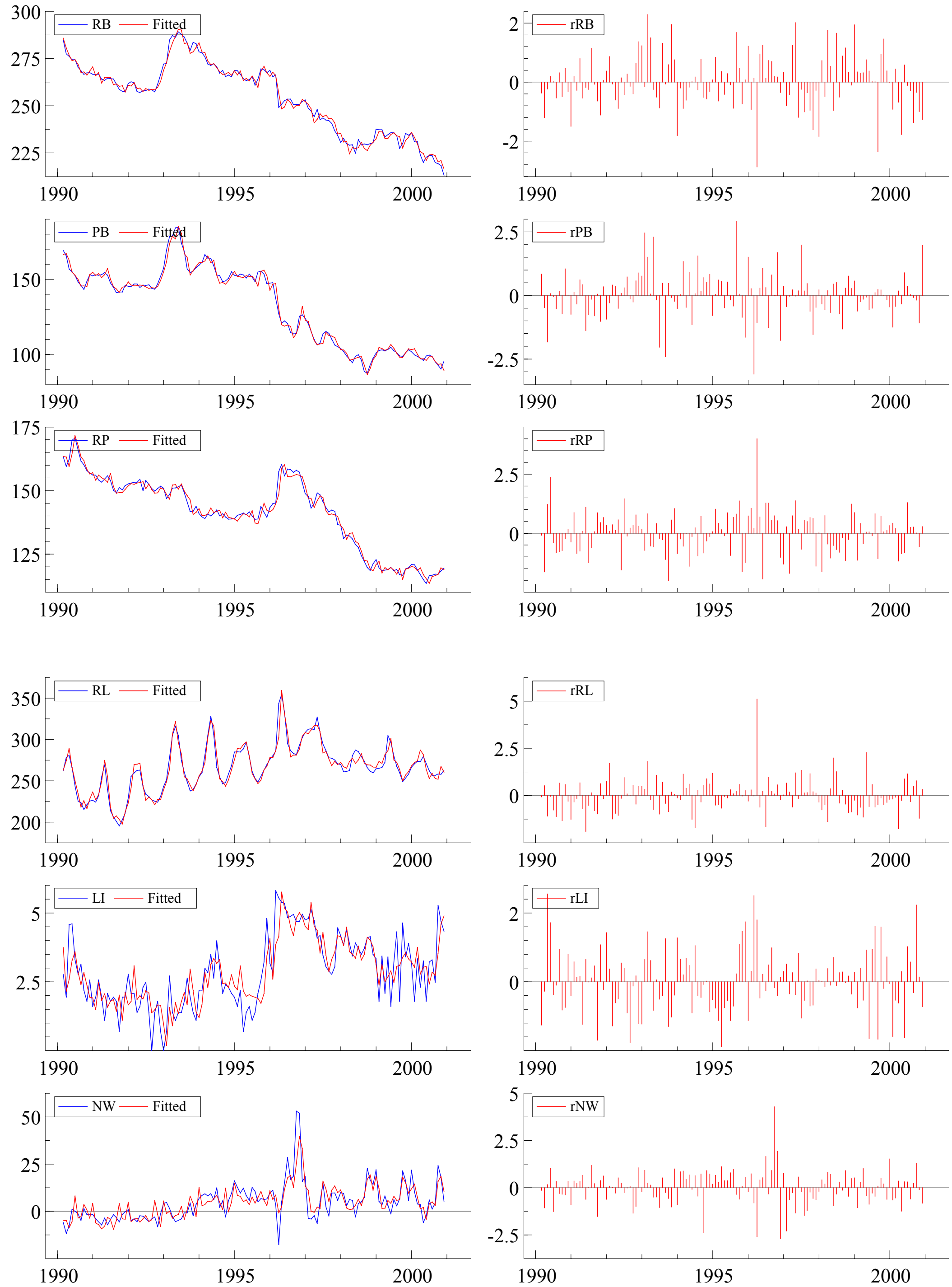
Smith, M.E., E.O. Raavensway and S.R. Thompson 'Sales Loss Determination in Food Contamination Incidents: An Application to Milk Bans in Hawaii' *American Journal of Agricultural Economics*, 70 (1988): 513-520.

Wohlgenant, M.K. ‘Marketing Margins: Empirical Analysis’ in G.C. Rausser and B.L. Gardner (eds.) *Handbook of Agricultural Economics*, Elsevier North-Holland, Elsevier Amsterdam.

Appendix : Plots of the VAR(2) model (unrestricted constant and seasonals)

(a) Actual and Fitted values

(b) Scaled Residuals



(c) Residuals from all potential co-integrating vectors

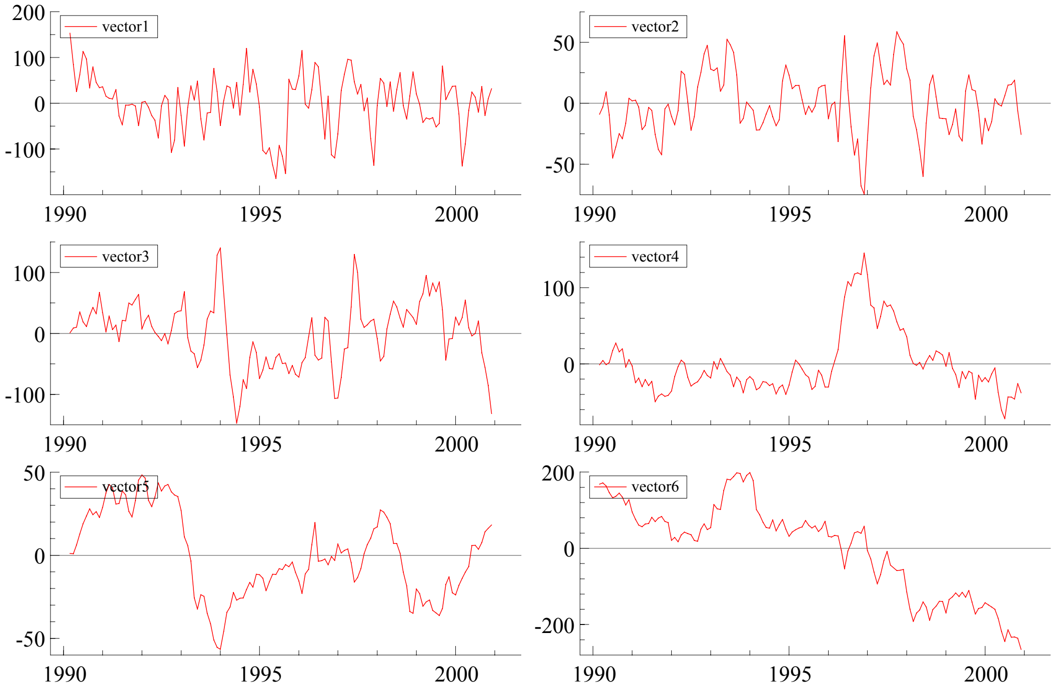
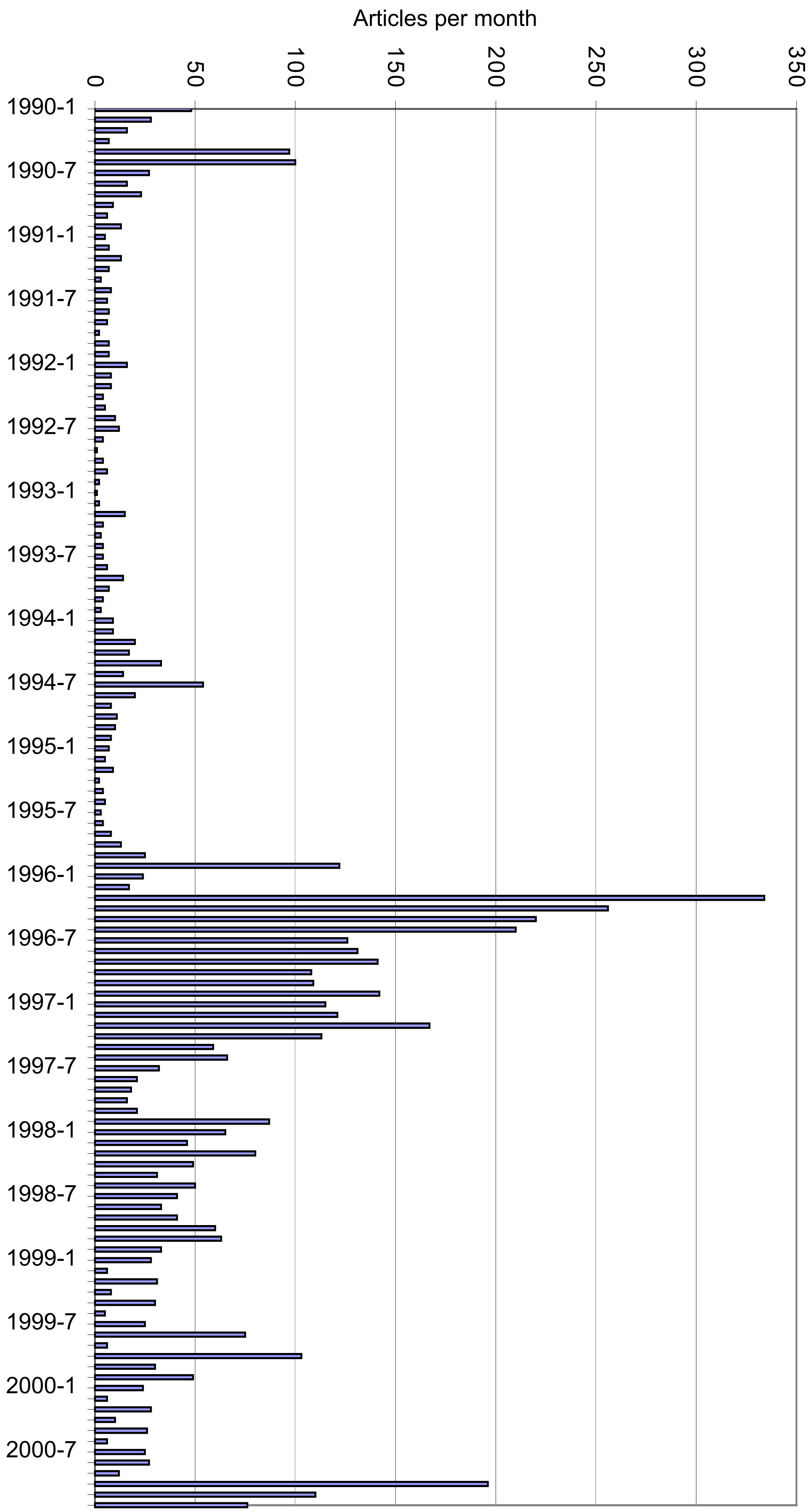


Figure 1: The Newspaper count (January 1990 to December 2000)



Articles per month
0
50
100
150
200
250
300
350
1990-1
1990-7
1991-1
1991-7
1992-1
1992-7
1993-1
1993-7
1994-1
1994-7
1995-1
1995-7
1996-1
1996-7
1997-1
1997-7
1998-1
1998-7
1999-1
1999-7
2000-1
2000-7

Figure 2 : Net Withdrawals from the UK Beef Sector (1990-2000)

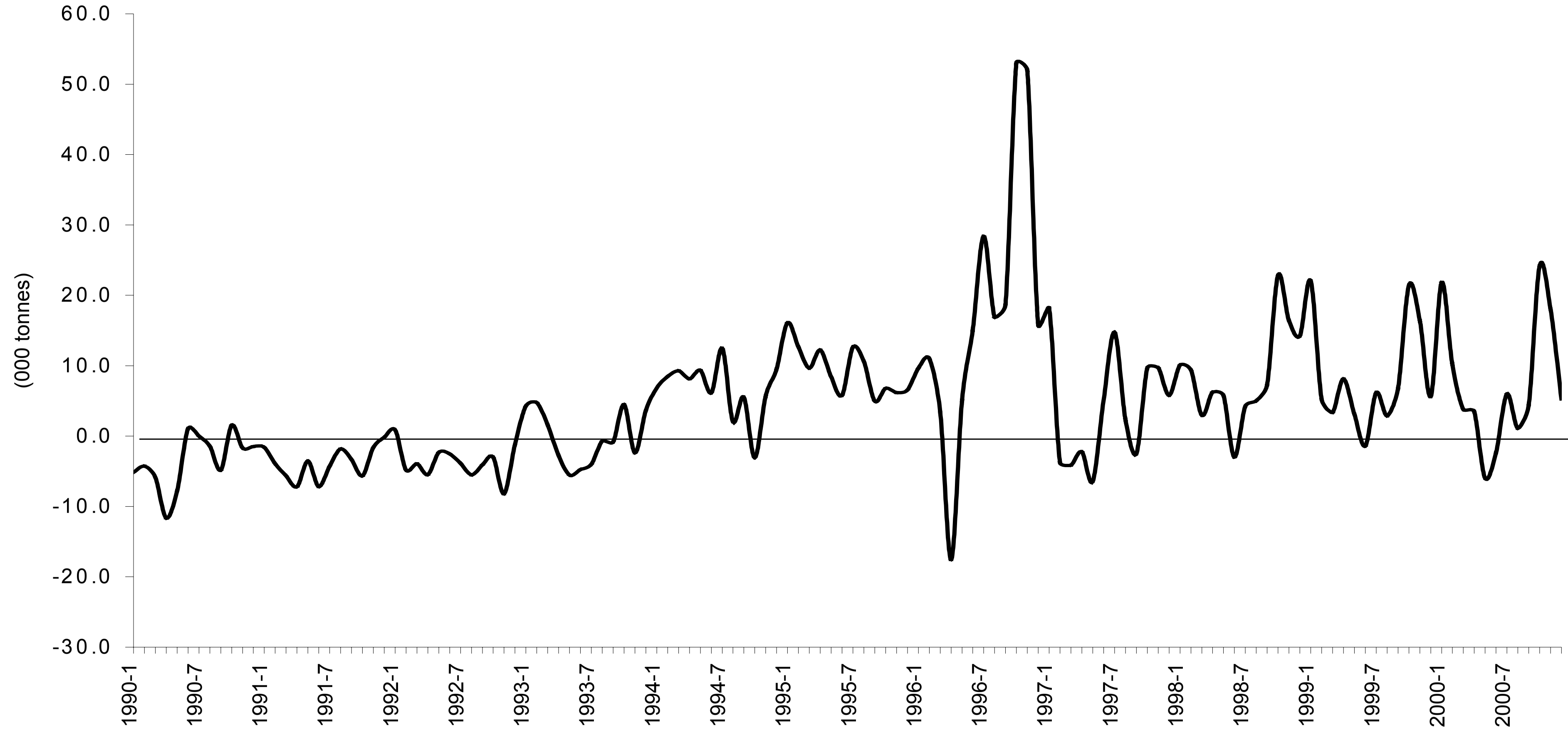

(000 tonnes)
60.0
50.0
40.0
30.0
20.0
10.0
0.0
-10.0
-20.0
-30.0
1990-1
1990-7
1991-1
1991-7
1992-1
1992-7
1993-1
1993-7
1994-1
1994-7
1995-1
1995-7
1996-1
1996-7
1997-1
1997-7
1998-1
1998-7
1999-1
1999-7
2000-1
2000-7

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

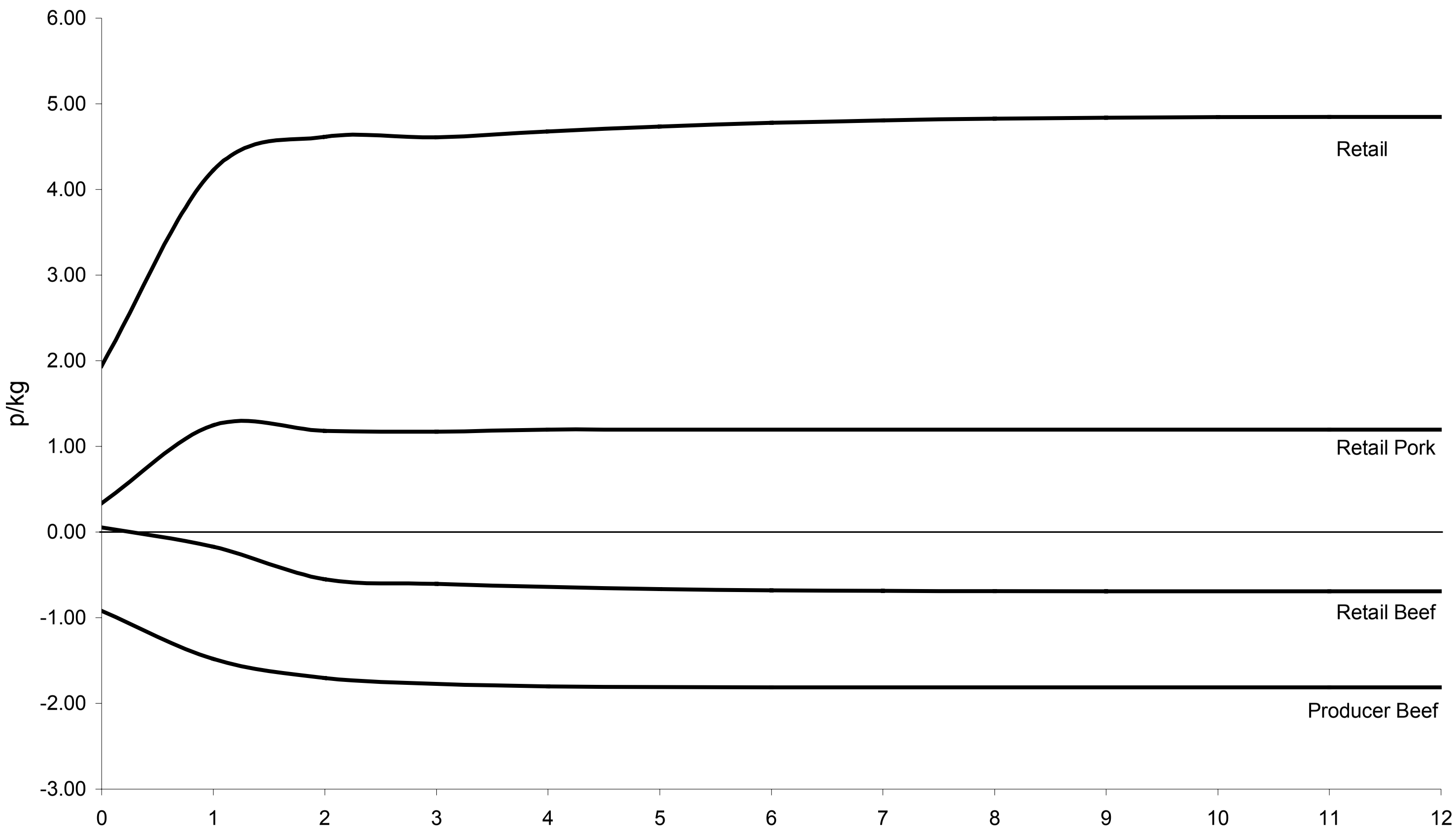


Figure 4: The Simulated Dynamic Effect of a (one standard error) Shock to Net Withdrawals