Price transmission in the Spanish bovine sector:

the BSE effect

Abstract

A regime-switching vector error correction model is applied to monthly price data to assess the impact of BSE outbreaks on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. To evaluate the degree to which price transmission is affected by BSE food scares, a BSE food scare index is developed and used to determine regime-switching. Results suggest that BSE scares affect beef producers and retailers differently. Consumer prices are found to be weakly exogenous and not found to react to BSE scares, while producer prices conversely adjusted. The magnitude of the adjustment is found to depend on the magnitude of the BSE scare.

Key words: Food scare, BSE crisis, price transmission, regime-switching.

1. Introduction

When suspicions arose that Bovine Spongiform Encephalopathy (BSE) was linked to a new-variant Creutzfeld-Jakob Disease (vCJD), a new food scare was generated. Apart from the obvious concerns about the impacts of the BSE on human health, there were also worries about the effect that the crisis might have on price transmission along the marketing chain. More specifically, there were suspicions that the crisis was having different impacts on retailers and producers. It was argued that producers were suffering stronger price declines than retailers, increasing the marketing margins in favor of retailers (Lloyd *et al.,* 2006) and threatening farmers' standard of living.

The objective of this paper is to formally assess the impacts of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. This analysis of the Spanish beef sector is important for several reasons. First, the magnitude of the BSE crisis in Spain has been significant. Spain ranks fifth in the European Union (EU)[1] regarding the number of BSE cases registered since 1987 (World Organization for Animal Health, 2007).

Second, the Spanish beef sector has substantial economic and social relevance both within Spain and within the EU. Beef is one of the most important activities in the Spanish agricultural sector, representing almost a fifth of the Spanish gross animal production (Ministerio de Medio Ambiente y Medio Rural y Marino, 2006). Also, Spain occupies the fifth position in the EU rankings for bovine production with 671 thousand tons, representing 9% of total output (Eurostat, 2007). The Spanish beef sector has relevance from a social perspective as well, since it constitutes, together with sheep farming, the main economic activity that can be carried out in several less-favored mountain areas. It is thus interesting to see how the BSE crisis affected the overall operation of this important economic sector.

This paper is organized as follows. In the second section, we present a brief overview of the Spanish BSE crisis. A literature review of previous research is presented in the third section. The fourth section is devoted to discussing econometric methods. The empirical application and the discussion of the results are presented in the fifth and sixth sections, respectively. Finally, the article ends with the concluding remarks section.

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[1] EU refers to EU-15.

2. The BSE in Spain

The disease Bovine Spongiform Encephalopathy was first identified in November 1986 in Britain by the government's Central Veterinary Laboratory at Weybridge. The number of BSE cases grew and spread both within the UK and to other EU countries. In particular, the first BSE cases identified outside Britain were in Ireland in 1989, Portugal and Switzerland in 1990, and France in 1991. In 1996 the UK Government confirmed the link between vCJD and BSE, confirming the most significant animal disease-related food scare in Europe (BBC NEWS, 2000).

In November 2000, the Spanish government announced the first two cases of BSE in Galicia, which is a large beef producing region, where more than one third of all BSE cases in Spain have been registered (Administración General del Estado, 2009). Several regulations intended to prevent the spread of the disease were passed. For instance, the Royal Decree 3454/2000 on December 22, 2000 was enforced to establish and regulate an integrated program to monitor and control the spread of the disease (Ministerio de la Presidencia, 2009). Regional authorities also struggled to manage the crisis with a shortage of veterinarians properly trained to identify the disease and only two national laboratories able to conduct the BSE tests ordered by the EU.

During the BSE crisis, Spanish beef production experienced significant changes. In particular, since no BSE cases were reported in Spain during the second half of the 1990s, production increased from 565 thousand tons, in 1996, to 678, in 1999. Production in 2000 was cut by almost 7% with respect to 1999, but recovered during the first half of the 2000s. Spanish foreign trade experienced even stronger changes. While Spanish beef exports increased from 1996 to 2000, they suffered a 21% decline in 2001. In 2002, the beef export market started to show signs of recovery. Imports increased during the second half of the 1990s but plummeted by 24% in 2001 and recovered in the subsequent years (Eurostat, 2007).

Spain has the fourth lowest per capita consumption of beef in the EU, above Belgium, Germany and Sweden (FAO, 2005). Before the BSE crisis, Spanish per capita consumption of beef remained relatively constant around 10 kg per capita per year. In 2001, beef consumption decreased by 19% to 7.9 kg per capita. Total beef consumption also declined by 18%. Consumption went back to normal levels after 2001 (Ministerio de Medio Ambiente y Medio Rural y Marino, 2005).

As noted above, Spain is fifth in the EU ranking of BSE cases registered since 1987, after the United Kingdom, Ireland, Portugal and France. The number of cows infected by BSE in Spain from 2000 to the end of 2007 was 717. As shown in figure 1, the number of cases increased from 82, in 2001, to a peak of 167 cases, in 2003. Four years later, the number had fallen to 36. This evolution of the BSE crisis in Spain is consistent with the whole EU situation where the number of BSE infected animals is also declining (World Organization for Animal Health, 2009).

3. Previous research: a literature review

Price is the primary mechanism through which different levels of the market are linked. The analysis of vertical price transmission allows for an approximation to the overall operation of the market (Goodwin and Holt, 1999). The magnitude of adjustment and speed with which shocks are transmitted are factors that reflect the competition in the food marketing chain and the market ability to adjust prices.

The analysis of vertical price transmission among various levels of the food marketing chain has gained special importance in the economics literature since the end of the last century. The attention devoted to these analyses has been partly motivated by social and political concerns that originated with the progressive concentration processes that occurred in the food industry and in the distribution sector. These changes may affect the competitive position of the economic agents participating in the market as well as price dynamics. Recent developments in time series econometrics have also contributed to the renewed interest in price transmission analyses.

Regime-switching models have received considerable attention within the price transmission literature. By making use of these techniques, several analyses have found that price shocks at one level of the marketing chain elicit different responses at other levels, depending on whether the shocks are positive or negative (see, for example, Goodwin and Piggott, 2001; Serra and Goodwin, 2003; Ben Kaabia and Gil, 2007). Another common finding of previous analyses is that prices at the retail level are sticky or slowly-responsive to price changes at other levels of the marketing chain.

Though a great number of studies and reports on the BSE crisis have been made available recently, literature regarding the economic impacts of food scares is relatively poor. Most of this literature has focused on the effects of food outbreaks on the demand for food (see, for example, Smith *et al.*, 1988; Burton and Young, 1996; Mazzocchi, 2006; Piggott and Marsh, 2004). A few have tried to assess whether the BSE crisis altered the efficiency with which the sector operates (Iraizoz *et al.*, 2005). However, the analyses of the effects of the BSE crisis on vertical price transmission are very scarce. A few notable exceptions are reviewed below.

Lloyd *et al.* (2001 and 2006) assess the impacts of the BSE crisis on vertical price transmission within the UK beef market. They use a cointegration framework that captures the relationship between beef prices and a food safety index reflecting the number of newspaper articles published on the topic. The authors assume that the food safety index is an exogenous shock to the system. Their results show that the BSE crisis has had a stronger impact on producer prices than on retail prices. More specifically, the effects on producer prices are more than double the impacts at the retail level.

Sanjuán and Dawson (2003) examine the impact of the BSE crisis on price transmission between producer and retailer prices for beef, lamb and pork in the UK. They use the cointegration procedure of Johansen *et al*. (2000) which allows for structural breaks in the cointegrated relationship. Their results show that a long-run relationship exists between producer and retail prices, which suffer from a structural break coinciding with the confirmation of the link between vCJD and BSE in 1996.

The unpublished analysis by Livanis and Moss (2005) studies the effects of food scares on price transmission in the U.S beef sector. Their methodological approach is similar to Lloyd *et al.* (2001), though they allow for structural breaks and consider the food index as endogenous. Their findings imply that information shocks are fairly transient in retail prices, but persist at the wholesale and farm levels. In another unpublished manuscript, Jaenicke and Reiter (2003) focus on analyzing the structural breaks that the BSE crisis may have caused to the cointegration relationship between producer and consumer beef prices in Germany. Their findings provide evidence that the BSE outbreak altered the producer-consumer price mechanism.

Common to the analyses revised above is the finding that the BSE crisis has altered price relationships and patterns of transmission along the beef marketing chain. The degree to which price transmission is affected by a food scare crisis is likely to depend on the scale of such crisis. Within this context, price relationships are likely to follow a regime-switching behavior and regime-switches are likely to be driven by the magnitude of the food scare. If markets respond differently to distinct levels of a food scare, marketing margins may change depending on the regime prevailing at each point in time. This article focuses on assessing this issue. To achieve the aforementioned objective, we estimate a regime-switching vector error correction model that captures the relationship between farm and retail beef prices. A BSE food scare index is developed and used as the variable determining regime-switching.

In spite of previous attempts to characterize price transmission responses to food scares, no previous analysis has allowed for regime-switching linked to the magnitude of the crisis, which represents a contribution of our work to the literature. Further, our analysis is the first to focus on the impacts of BSE on price transmission within the Spanish beef market.

4. Econometric methods

4.1. Regime-Switching Autoregressive Models

As previously mentioned, many empirical analyses have found evidence of nonlinear price transmission in food markets. Several econometric techniques have been developed to capture these nonlinearities. Early analyses were based on the econometric model introduced by Wolffram (1971) and later refined by Houck (1977) and Ward (1982). These specifications have been criticized for ignoring the time-series properties of the (usually nonstationary) data. Von-Cramon Taubadel (1998) extended the Wolffram specification to allow for an error correction term. More recently, Goodwin and Holt (1999) proposed using threshold vector error correction models (TVECM) to allow error correction specifications to adequately capture nonlinear and threshold-type price adjustments.

Tong (1978) originally introduced nonlinear threshold time series models. Tsay (1989) developed a methodology to test for threshold nonlinearity based on autoregressions and to model threshold autoregressive processes. Balke and Fomby (1997) combined error correction models and the threshold autoregressive model developed by Tsay (1989) into a threshold error correction framework. These authors suggest implementing a grid search procedure to select the threshold parameters that delineate different regimes through minimizing the sum of squared errors.

Consider a standard linear cointegration relationship between two variables, $P_{1,t} - \beta P_{2,t} = v_t$, where $P_{1,t}$ and $P_{2,t}$ are prices at different levels of the marketing chain and v_t represents the deviation from the equilibrium relationship. As is well known, cointegration between the two price series depends on the nature of the autoregressive process $\Delta v_t = \gamma v_{t-1} + u_t$. A value of gamma close to one implies that deviations from the equilibrium are stationary and that the price series are cointegrated.

Following Balke and Fomby (1997), this analysis can be extended to a regime-switching autoregressive (RSAR) process. A two-regime RSAR can be expressed as:

$$\Delta v_t = \begin{cases} \gamma^{(1)} v_{t-1} & if \quad s_{t-d} \le c \\ \gamma^{(2)} v_{t-1} & if \quad s_{t-d} > c \end{cases} \tag{1}$$

where s_{t-d} is the variable determining regime-switching and d is the lag of this variable (in this application we assume a delay of $d = 1$). Parameter c represents the threshold that delineates the different regimes and $\gamma^{(m)}, m = 1$ and 2 are speed of adjustment parameters that measure the rate at which prices adjust to disequilibria from the long run equilibrium relationship. The RSAR can be alternatively expressed as $\Delta v_t = \gamma^{(1)} v_{t-1} d_{1t}(c,d) + \gamma^{(2)} v_{t-1} d_{2t}(c,d) + u_t$, where $d_{1t}(c,d) = 1(s_{t-d} \le c)$ and $d_{2t}(c,d) = 1(s_{t-d} > c)$. When the threshold variable is a lagged residual of the error correction term, $s_{t-d} = v_{t-d}$, the model is known as a threshold autoregressive (TAR) model.

4.2. Regime-Switching Vector Error Correction Models

A Regime-Switching Vector Error Correction Model (RSVECM) is a multivariate version of the RSAR model. The RSVECM allows one to uncover potential nonlinearities in the adjustment of individual prices and provides more information about short-run price dynamics. Lo and Zivot (1999) have suggested that the multivariate RSVECM has higher power than univariate RSAR models. RSVECMs occur when some forcing variable (the variable relevant to the threshold behavior) leads to switching between different regimes and the variables in the model exhibit different types of behavior in each regime. Different regimes are represented by different parameter estimates of the underlying model. A two-regime RSVECM can be expressed as follows:

$$\Delta P_t = \begin{cases} \alpha^{(1)} + \alpha_p^{(1)} v_{t-1} + \sum_{i=1}^{l} \alpha_i^{(1)} \Delta P_{t-i} + \varepsilon_t^{(1)} & if \quad s_{t-d} \leq c \\ \alpha^{(2)} + \alpha_p^{(2)} v_{t-1} + \sum_{i=1}^{l} \alpha_i^{(2)} \Delta P_{t-i} + \varepsilon_t^{(2)} & if \quad s_{t-d} > c \end{cases} \tag{2}$$

where $P_t = (P_{1t}\ P_{2t})$ is the vector of prices being analyzed, $\alpha^{(m)}, \alpha_i^{(m)},\ m = 1$ and 2 are parameters showing the short-run dynamics and $\alpha_p^{(m)}$ are the speed of adjustment parameters that measure the speed at which the adjustment of prices to deviations from the long-run equilibrium takes place.[2] The RSVECM can be compactly expressed as:

$$\Delta P_t = \begin{cases} \boldsymbol{\alpha}^{(1)'} x_{t\text{-}1} + \varepsilon_t^{(1)} & if \quad s_{t-d} \leq c \\ \boldsymbol{\alpha}^{(2)'} x_{t\text{-}1} + \varepsilon_t^{(2)} & if \quad s_{t-d} > c \end{cases} \tag{3}$$

where:

[2] When $s_{t-d} = v_{t-d}$, the RSVECM is known as TVECM. TVECM have been used by Goodwin and Holt (1999), Goodwin and Harper (2000), Goodwin and Piggott (2001), Goodwin *et al.* (2002), Serra and Goodwin (2003), Serra and Goodwin (2004), or Ben Kaabia and Gil (2007).

$$x_{t-1} = \begin{bmatrix} 1 \\ v_{t-1} \\ \Delta P_{t-1} \\ \vdots \\ \Delta P_{t-l} \end{bmatrix}$$

and $\alpha^{(m)}$, is a vector of parameters.

Our specific estimation strategy can be summarized as follows. First, standard unit root and cointegration tests are conducted in order to determine whether price series are stationary and whether they are cointegrated, respectively. Based on the assumption that different levels of food scares can lead to different price adjustment, we next estimate a two-regime RSVECM where a food scare information index is used as the threshold variable.[3] Details on how this information index is built are provided below in subsection 4.4. Finally, we utilize the sup-LR statistic developed by Hansen and Seo (2002) to test for a linear VECM against the alternative of a RSVECM. Details on specification tests are given in the next subsection.

The parameters of the multivariate RSVECM can be estimated using sequential multivariate least squares in two steps. In the first step, a grid search is carried out to estimate the threshold parameter, c.[4] The threshold is searched over the values of the threshold variable and the search is restricted to ensure an adequate number of observations for estimating the parameters in each regime. Recently, Serra and Goodwin (2004) considered two general grid search approaches in the selection of thresholds which may not be equivalent. The first approach involves minimization of

[3] The variable relevant to threshold behavior has been usually assumed to be the (lagged) error correction term and the threshold has been usually searched over the values of this term (see, for example, Goodwin and Piggott, 2001; or Serra and Goodwin, 2003).

[4] Several analyses that are based on threshold models have treated regime-switching as exogenous (see, for example, Chavas and Mehta, 2004). We adopt a more general model that incorporates this issue as endogenous.

the sum of squared errors or, alternatively, the trace of the covariance matrix of the residual errors. This approach has been implemented by a number of empirical analyses (see, Balke and Fomby, 1997; or Goodwin and Piggott, 2001). The second approach maximizes a likelihood function (see, for example, Obstfeld and Taylor, 1997; or Moschini and Meilke, 1989). As Serra and Goodwin (2004) explain, the kernel of the likelihood function involves the logged determinant of the residual covariance matrix. In our analysis, we follow the latter approach because, contrary to the first alternative, it does not ignore the cross equation correlation.

Under the specified approach with variance-covariance matrix (Σ), the vectors of parameters $\boldsymbol{\alpha}^{(m)}, m=1$ and 2 are estimated by iterated seemingly unrelated regressions (SUR) method giving:

$$S(c,d) = 1n\left|\hat{\Sigma}(c,d)\right|, \tag{4}$$

where $\hat{\Sigma}(c,d)$ is a multivariate SUR estimate of $\Sigma = \mathrm{var}(\varepsilon_t)$, conditional on (c,d) where $d=1$. The vector of the errors is represented by $\varepsilon_t' = \left(\varepsilon_t^{(1)} \quad \varepsilon_t^{(2)}\right)$. In the second step, the least squares estimates of c are obtained based on:

$$(\hat{c},1) = \arg \min_{c} S(c,1). \tag{5}$$

Final parameter estimates are given by $\hat{\boldsymbol{\alpha}}^{(m)} = \hat{\boldsymbol{\alpha}}^{(m)}(\hat{c},1)$ and the estimations of the residual covariance matrix by $\hat{\Sigma} = \hat{\Sigma}(\hat{c},1)$.

4.3. Specification tests

After estimating the model, we test for the significance of the differences in parameters across relative regimes. The sup-LR test developed by Hansen and Seo (2002) is used to test for a linear VECM against the alternative of a RSVECM. The model under the null is $\Delta P_t = \boldsymbol{\alpha}' x_{t-1} + \varepsilon_t$, while the model under the alternative can be expressed as $\Delta P_t = \boldsymbol{\alpha}^{(1)'} x_{t-1} d_{1t}(c,d) + \boldsymbol{\alpha}^{(2)'} x_{t-1} d_{2t}(c,d) + \varepsilon_t$. The sup-LR statistic can be computed in the following way:

$$LR = T\left(\ln\left|\hat{\Sigma}\right| - \ln\left|\hat{\Sigma}\right|(\hat{c},d)\right), \qquad (6)$$

where $\hat{\Sigma}$ is the variance-covariance matrix of the residuals for the VECM, $\hat{\Sigma}(\hat{c},d)$ represents the variance-covariance matrix of the residuals of the RSVECM and T is the number of observations. The sup-LR statistic has a non-standard distribution because the threshold parameters are not identified under the null hypothesis. To calculate the p-value of the sup-LR statistic we carry out the residual bootstrap technique developed by Hansen and Seo (2002). A total of 500 simulations are run.

More specifically, we carry out a parametric residual bootstrap method to approximate the sampling distribution that requires the specification and estimation of the model under the null (VECM), an assumption on the distribution of the residuals of the model and the initial conditions. We assume that model residuals are distributed as a Normal $(0,\Sigma)$, where Σ is the covariance structure of the original VECM. The initial values of the model variables are set equal to their actual values. Given the initial conditions, we generate random shocks of the residuals at each period and derive the vector of series by recursion. The sup-LR test is computed for each replicated sample and stored. The bootstrap p-value can be determined as the proportion of simulations under the null for which the simulated LR statistic exceeds the observed statistic.

4.4. The food scare information index

Our analysis evaluates the impacts of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. As noted above, some evidence has suggested that the BSE crisis had an impact on the vertical price transmission processes (Lloyd *et al.*, 2001, 2006). The degree to which price transmission is affected by a food scare crisis is likely to depend on the scale of the crisis. A food scare information index (FSII), based on the number of news articles published on the topic, is built in order to have a measure of this scale. The FSII is used as a threshold variable. Hence, we implicitly assume that different levels of food scare can lead to different price reactions and that the food scare index leads to switching between different price behavior regimes.

The literature has suggested various methods to construct a FSII based on the news count. In our analysis we use the method proposed by Chern and Zuo (1997) to build a monthly food scare information index. Specifically, Chern and Zuo (1997) extended the cumulative method used by Brown and Schrader (1990) by building a new fat and cholesterol information index that considers a differentiated carryover weight for supporting and non-supporting articles. Also, the articles are assumed to have a finite duration and lag distribution as a source of information. The FSII index based on this method can be expressed as:

$$FSII_t = \sum_{i=0}^{n} W_i \ \mathrm{NM}_{t-i} \tag{7}$$

where NM_{t-i} is the number of relevant articles (both supporting and non-supporting) published during period *t-i*, W_i is the weight attributed to the lagged period i and *n* is number of lagged periods considered. This method not only allows for a carryover effect but also for a decay effect of information. The carryover and decay effects are captured by specifying the weight function and the total lag period. Chern and Zuo (1997) utilize a cubic or third-degree polynomial weight function (CWF) because it generates asymmetric weights. The cubic weight function can be written as:

$$W_i = \lambda_0 + \lambda_1 i + \lambda_2 i^2 + \lambda_3 i^3 \tag{8}$$

where the λs are parameters and i is the number of lagged periods. The values of the coefficients need to be determined based on the following restrictions. First, the maximum weight lies somewhere between the current period ($i = 0$) and the last period ($i = n$). Second, the minimum weight occurs at $i = n + 1$ and is set to zero $W_{n+1} = 0$. Finally, the sum of weights over the current and lagged periods is equal to 1 $\left(\sum_{i=0}^{n} W_i = 1 \right)$. The cubic weight function can be rewritten as (see Chern and Zuo, 1997):

$$W_i = 2a/((n+1)b) + (12m/b)i - (6(n+1+m)/((n+1)b))i^2 + (4/((n+1)b))i^3 \tag{9}$$

where $a = (n+1)^2 (n+1-3m)$ and $b = (n+2)[(n+1)^2 - m(2n+3)]$. The lag period with the maximum weight is represented by *m*. Expression $(n+1-3m)$ is restricted to be positive. Both *n* and *m* can take any finite number.

5. Empirical application

Our empirical analysis utilizes two series of monthly beef prices and one monthly series representing the food scare information index. Beef prices are observed from January 1996 to December 2005, giving a total of 120 observations. Price information includes both farm-gate prices for prime beef (1 to 2 years old) expressed in euros per 100 kilo and consumer prices expressed in euros per kilo. Both prices were obtained from the Ministerio de Medio Ambiente y Medio Rural y Marino.[5]

As noted above, health risks have received increasing attention among consumers in developed countries and strengthened the relationship between food scares, food consumption, and prices. More specifically, if consumers believe that beef is unsafe to eat, there will be a decline in the demand and possibly the price of beef. To investigate the impacts of food scare concerns during the BSE crisis on price relationships and patterns of vertical price transmission within the Spanish bovine sector and following the methodology outlined in subsection 4.4, we construct a FSII that captures the degree of the consumer food scare. The index is based on a monthly count of newspaper articles on the BSE crisis appearing in a major Spanish newspaper, EL PAÍS.

Weinberger and Dillon (1980) suggest that supporting articles (unfavorable news) may be more influential than a similar amount of non-supporting articles (favorable news). It has been also reported that a similar quantity of unfavorable news weighs far more heavily on consumer decision-making than favorable news (Chang and Kinnucan, 1991; Kinnucan and Myrland, 2000). However, Mazzochi (2006) argues that discrimination between favorable and unfavorable articles is highly

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[5] Farm gate prices are obtained from Boletín Mensual de Estadística, while consumer prices were made available to the authors upon request by the Subdirección General de Industria, Innovación y Comercialización Agraria. In the latter case, a weighted average of retail prices from the main cuts sold in the market is used.

subjective. This can be especially true with the BSE crisis, since the long latent period of vCJD will not cause the same impact on the young than on the old population. Due to the aforementioned reasons and following Kim and Chern (1997) and Chern and Zuo (1997), we do not weigh articles depending on whether they are favorable or unfavorable.

We construct our monthly FSII over the period from July 1995 to December 2005, by scanning all articles relevant to BSE crisis using different keywords.[6] The keywords searched were "Vaca(s) Loca(s)" (mad cow), "Encefalopatía Espongiforme Bovina" (Bovine Spongiform Encephalopathy), and "Creutzfeldt-Jakob." The number of news articles ranged from a maximum of 354 in February 2001, to a minimum of 0 in a few months far from the peak of the crisis. The average was 20.8 news articles per month with a standard deviation of 44.12. By excluding geographical indicators from the keyword list, our FSII intends to include both the information on the development of the crisis in Spain and in other EU countries. Due to the important trade relationships between Spain and the EU, Spanish markets are likely to have been affected by the development of the food scare in other EU countries. As a result, despite the fact that the BSE was first identified in November 2000 in Spain, the FSII is developed from 1995, in order to capture the impacts of earlier infections in other countries.

Following Chern and Zuo (1997), we computed our monthly FSII using different values for both the number of lags (n) and peak times (m) and found the index to remain relatively stable across these values. In light of these results, we selected n=6 which is consistent with the

[6] Though our analysis focuses on the January 1996 to December 2005 period, the monthly count was carried out for a longer period in order to allow for lags in the computation of the FSII.

recommendation by Clarke (1976) and requires m to be equal or less than two. Since the FSII is very similar independently of the chosen value for the peak time, we selected m=1.[7]

The monthly food scare information index used in our analysis is presented in figure 2. A comparison of this figure with figure 1 containing the evolution of the BSE cases in Spain suggests that the perception that economic agents have on the crisis does not necessarily keep pace with the number of infected animals. Small index increases occurring around 1996 are due to mass media reporting the link between the vCJD and the BSE. The index skyrockets during 2000-2001 when the BSE finally entered Spain. Additionally, after 2002 the BSE cases and the FSII follow different paths. In spite of the fact that infections continued to be found after the onset of the crisis, they were reported with less emphasis by mass media.

A dummy variable representing the months when the crisis peaked in terms of mass media reporting (from February to April 2001) is introduced as a regime-independent variable in the short-run specification of the RSVECM. This dummy also coincides with the first BSE cases detected in Spain.

6. Results

As noted above, our empirical analysis utilizes monthly farm-gate and consumer beef prices. It is also based upon a newspaper index that is a measure of the degree of food scare on a monthly basis. The empirical analysis is based on a logarithmic transformation of prices. Standard Dickey

[7] Results of the index using different lags and peak times are not presented here, but are available from the authors upon request.

and Fuller (1979) and KPSS (Kwiatkowski *et al.*, 1992) tests for each price series provide evidence that all price series are integrated of order one I(1).[8]

After testing for unit roots, the Engle and Granger (1987) test for cointegration is used. In doing so, the equilibrium relationship is normalized by the producer price and ordinary least squares (OLS) are used to obtain estimates of the cointegrating parameters (see table 1). The normalization variable is selected according to previous research results that confirm that while producer prices tend to adjust to their long-run parity, consumer prices are sticky or slowly-responsive to price changes occurring at other levels of the marketing chain (see Goodwin and Holt, 1999; Borenstein *et al.*, 1997; Peltzman, 2000). This is confirmed later in our analysis. The Engle and Granger (1987) test indicates the existence of a cointegration relationship among producer and consumer price series (see table 2).[9] Other analyses have also found evidence of cointegration between producer and consumer prices for beef (Lloyd *et al.*, 2001; Sanjuán and Dawson, 2003; Jaenicke and Reiter, 2003; Livanis and Moss, 2005).

[8] The Akaike Information Criterion (AIC) as well as Schwartz Bayesian Criterion (SBC) are used to select the proper lag length of the autoregressive process. When the two criteria differ, we use the more parsimonious SBC criteria (Enders, 1995, p. 88; Wang and Liu, 2006). Results are available from the authors upon request.

[9] It is important to note that Johansen (1988) cointegration test is also applied and suggests that the hypothesis of no cointegration can be rejected at the 5% significance level. Furthermore, results present no significant differences when using the error correction term obtained from Johansen or Engle and Granger method. In light of these results, we select the latter test which is consistent with the recommendation by Enders (1995, p.385) in the presence of a single cointegration vector. The LM test for constancy of the cointegration parameters (Hansen and Johansen, 1999) suggests constancy of these parameters throughout the period studied. Results are available from the authors upon request.

A RSVECM is then estimated by using sequential multivariate least squares in two steps[10]. The threshold derived from the grid search and the sup-LR statistic are presented in table 3.[11] The sup-LR test statistic indicates that nonlinearities are statistically significant at the 95% confidence level.[12] This involves the existence of different price behavior depending on the number of news articles published on the BSE crisis. Specifically, price behavior can be classified into two regimes, one corresponding to food scare index values FSII<=31.218, and a second regime corresponding to FSII>31.218. In table 4 we present the RSVECM parameter estimates across the different regimes.

Parameter estimates suggest that producer prices adjust to deviations from the long-run equilibrium relationship. Conversely, consumer prices are exogenous and do not adjust to system disequilibria. The BSE crisis seems to have affected Spanish beef farmers and retailers differently, since all adjustments to deviations from equilibrium occur at the producer level. The producer price presents two different adjustments. More specifically, producer price responses are slower (-37%) when the weighted number of news articles is not greater than 31 (regime I), than when the news are above this threshold (-70% for regime II). The response in the latter regime is almost double the response in the first regime.

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[10] The optimum number of lags is selected by considering the Breusch-Godfrey Lagrange Multiplier test for autocorrelation and the SBC and AIC criterion. Though the SBC and AIC criterion recommend the use of just one lag, three lags are used in the estimation to avoid autocorrelated residuals.

[11] The bootstrap process needed to compute the p-value of the sup-LR test is computationally intensive. To keep computations manageable, we limit the number of simulations to 500 for each model.

[12] A three regime model was tested against a two regime one and results suggest no significant differences between the two models. Results are not presented here, but are available from the authors upon request.

Figure 3, presents the timing of jumps between the two alternative regimes.[13] The figure illustrates that regime I, which has the majority of observations, coincides with periods following an important crisis, i.e. the crisis following the confirmation in 1996 by the UK government that the BSE is linked to the vCJD, and the crisis originated by the contamination of Spanish cattle at the end of the 2000s. These two crises are represented by regime II (threshold higher than 31), where the strongest adjustment takes place.

While our results are compatible with previous research results on the BSE crisis, the use of a RSVECM offers an advantage over previous studies based on cointegration relationships, since it allows for different adjustment processes depending on the market situation. In this regard, our analysis represents a contribution to previous studies on the impacts of the BSE crisis on price transmission mechanisms. In using the FSII as the threshold variable, we are able to assess how price adjustment changes depending on the degree of food scare.

7. Concluding remarks

Food borne diseases are not only a relevant public health issue, but they also have important economic implications. The objective of this paper is to formally assess the impacts of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. The degree to which price transmission is affected by a food scare crisis is likely to depend on the scale of such crisis. To the extent that this occurs, price relationships are likely to

[13] It is important to note here that the behavior of the threshold variable is parallel to the behavior of the error correction term. The error correction term reaches its peaks when the food scare information index is highest and is closer to zero for low levels of the index. This is expected and allows forecasting that the use of the error correction term instead of the index as the threshold variable may not have substantially altered the results.

show a threshold-type behavior and regime-switching is likely to be driven by the degree of food scare. To capture this issue a food scare information index is used.

In spite of previous attempts to characterize price transmission responses to food scares, no previous analysis has allowed for regime-switching linked to the magnitude of the crisis, which represents a contribution of our work to the literature. Our analysis also contributes to previous literature in that we focus on the impacts of BSE on price transmission within the Spanish beef market, a market that has not been investigated.

To achieve the aforementioned objective, we estimate a RSVECM. Our empirical model utilizes two series of monthly farm and retail beef prices and one monthly series representing the food scare information index. The food scare information index is built following Chern and Zuo (1997) and based on a count of newspaper articles on the zoonosis that appeared in a major Spanish newspaper.

The results of this paper can be summarized as follows. Standard unit root tests confirm the presence of a unit root in each price series. Cointegration tests provide evidence of a long-run equilibrium relationship between producer and consumer prices. Other analyses have also found a long-run equilibrium relationship between beef prices at different levels of the marketing chain.

The estimated RSVECM suggests that the BSE crisis affects beef producers and retailers differently. Consumer prices, which are found to be exogenous, do not adjust to disequilibria caused by the food scare. Conversely, producer prices are endogenous and do all the adjustment. These results are expected and are compatible with previous research that has suggested that upstream prices in the marketing chain generally do all the adjustment, while consumer prices are sticky and slowly-responsive. Finally, the sup-LR test statistic indicates that nonlinearities are statistically significant.

If only producer prices adjust, producer margins will be squeezed while retail margins will not. Distributional issues will thus arise. This will increase the likelihood of producers abandoning the sector during strong crises and probably increases the need for public intervention if this is to be prevented. Reinforcing regulatory compliance as well as private food safety investments should be useful in this regard. Strengthening food safety measures from the farm to the table would increase reputation for safe food production, and reduce contamination of food and food born diseases. The costs these measures have on small producers may pay off the costs they would face in case of a food scare developed due to a lack of controls. This however remains an area for further research.

This paper can be extended in a number of ways. First, it would be useful to implement the proposed methodology to other food scare crises and see if the same conclusions hold. Serra *et al.* (2006) explained that non-parametric techniques can be used to overcome the limitations involved with parametric threshold models. This would allow for an evaluation into the extent to which our results are subject to the specific functional forms used in the analysis.

References

Administración General del Estado, 2009. Información EEB. Accessed March 2009, available at http://www.eeb.es/ .

Balke, N. S., Fomby, T. B., 1997. Threshold cointegration. *Int. Econ. Rev.* **38**, 627-645.

BBC NEWS, 2000. BSE and CJD: Crisis and Chronology. Accessed March 2009, available at http://news.bbc.co.uk/hi/english/static/in_depth/health/2000/bse/1986.stm.

Ben Kaabia, M., Gil, J. M., 2007. Asymmetric price transmission in the Spanish lamb sector. *Europ. Rev. Agr. Econ.* **34**, 53-80.

Borenstein, S., Cameron, A., Gilbert, R., 1997. Do gasoline prices respond asymmetrically to crude oil prices?. *Quart. J. Econ.* **112**, 305-339.

Brown, D. J., Schrader, I. F., 1990. Information on cholesterol and falling shell egg consumption. *Amer. J. Agr. Econ.* **73**, 548-555.

Burton, M., Young, T., 1996. The impact of BSE on the demand for beef and other meats in Great Britain. *Appl. Econ.* **28**, 687-693.

Chang, H. S., Kinnucan, H., 1991. Advertising, information and product quality: the case of butter. *Amer. J. Agr. Econ.* **73**, 1195-1203.

Chavas, J. P., Mehta, A., 2004. Price dynamics in a vertical sector: the case of butter. *Amer. Agr. Econ. Assoc.* **86**, 1078-1093.

Chern, W. S., Zuo, J., 1997. Impacts of fat and cholesterol information on consumer demand: Application of news indexes. Working paper. The Ohio State University, Columbus, Ohio.

Clarke, D., 1976. Econometric measurement of the duration of advertising effects on sales. *J. Marketing Res.* **13**, 345-357.

Dickey, D. A., Fuller, W. A., 1979. Distribution of the estimators for autoregressive time Series with a unit root. *J. Amer. Statistical Assoc.* **74**, 427-431.

Enders, W., 1995. *Applied Econometric Time Series*. Iowa State University. Johan Wiley & Sons, INC.

Engle, R. F., Granger, C. W. J., 1987. Co-integration and error correction: representation, estimation and testing. *Econometrica* **55**, 251-276.

Engle, R. F., Yoo, B. S., 1987. Forcasting and testing in cointegrated systems. *J. Pol. Modeling* **35**, 143-159.

Eurostat, 2007. Dataset. Accessed July 2007, available at europa.eu.int/comm/eurostat/.

FAO, 2005. Dataset. Accessed July 2007, available at http://faostat.fao.org/default.aspx.

Goodwin, B. K., Holt, M., 1999. Price transmission and asymmetric adjustment in the U.S. beef sector. *Amer. J. Agr. Econ.* **81**, 630-637.

Goodwin, B. K., Harper, D., 2000. Price transmission, threshold behavior, and asymmetric adjustment in the US pork sector. *J. Agr. Appl. Econ.* **32**, 543-553.

Goodwin, B. K., Piggott, N. E., 2001. Spatial marketing integration in the presence of threshold effects. *Amer. J. Agr. Econ.* **83**, 302-317.

Goodwin, B. K., Grennes, T., Craig, L., 2002. Mechanical refrigeration and the integration of perishable commodity market. *Exploration Econ. Hist.* **39**, 154-182.

Hansen, H., Johansen, S., 1999. Some tests for parameter constancy in cointegrated VAR-models. *Econometrics J.* **2**, 306-333.

Hansen, B. E., Seo, B., 2002. Testing for two-regime threshold cointegration in vector error correction model. *J. Econometrics* **110**, 293-318.

Houck, P. J., 1977. An approach to specifying and estimating non reversible functions. *Amer. J. Agr. Econ.* **59**, 570-572.

Iraizoz, B., Bardaji, I., Rapun, M., 2005. The Spanish beef sector in the 1990s: impact of the BSE crisis on efficiency and profitability. *Appl. Econ.* **37**, 473-484.

Jaenicke, J., Reiter, I., 2003. Development of prices during the BSE crisis. Some empirical and theoretical results. Paper presented at the 4th Workshop for Macroeconomtrics, Halle Institute for Economic Research, University of Osnabrück, Germany.

Johansen, S., 1988. Statistical analysis of cointegration vectors. *J. Econ. Dynam. Control* **2**, 424-438.

Johansen, S., Mosconi, R., Nielsen, B., 2000. Cointegration analysis in the presence of structural breaks in the deterministic trend. *Econometrics J.* **3**, 216-249.

Kim, S. R., Chern, W. S., 1997. Indicative measures of health risk information on fat and cholesterol for U.S. and Japanese consumers. *Consumer Interests Annual* **43**, 84-89.

Kinnucan, H. W., Myrland, O., 2000. Relative impact of health information and advertising on commodity markets. In: Proceedings of the XXIV IAAE Conference, Mini Symposium "Effects of Health Information on the Demand for Food," Berlin, Germany.

Kwiatkowski, D., Phillips, C. B., Schmidt, P., Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?. *J. Econometrics* **54**, 159-178.

Livanis, G., Moss, C. B., 2005. Price transmission and food scares in the U.S. beef sector. Paper presented at the American Agricultural Economics Association Meetings in Providence, Rhode Island.

Lloyd, T., McCorriston, S., Morgan, C. W., Rayner, A. J., 2001. The impact of food scares on price adjustment in the UK beef market. *Agr. Econ.* **25**, 347-357.

Lloyd, T. A., McCorriston, S., Morgan, C. W., Rayner, A. J., 2006. Food scares, market and price transmission: the UK BSE crisis. *Europ. Rev. Agr. Econ.* **33**, 119-147.

Lo, M. C., Zivot, E., 1999. Threshold cointegration and nonlinear adjustment to the law of one price. University of Washington.

Mazzocchi, M., 2006. No news is good news: Stochastic parameters versus media converage indices in demand models after food scares. *Amer. J. Agr. Econ.* **88**, 727-741.

Ministerio de la Presidencia, 2009. Disposiciones generales. Accessed February 2009, available at http://www.boe.es/boe/dias/2000/12/23/pdfs/A45550-45565.pdf.

Ministerio de Medio Ambiente y Medio Rural y Marino, 2005. Sacrificio Total de Ganado Bovino. Accessed May 2007, available at http://www.mapa.es/estadistica/pags/sacrificio/bovino2005.xls.

Ministerio de Medio Ambiente y Medio Rural y Marino, 2006. Evolución Macromagnitudes Agrarias 1990-2006. Accessed March 2009, available at http://www.mapa.es/estadistica/pags/macromagnitudes/CEA_Resultados_Nacionales_1990-2006_enero_2007.xls .

Moschini, G., Meilke, K. D., 1989. Modeling the pattern of structural change in U.S. meat demand. *Amer. J. Agr. Econ.* **71**, 253-262.

Obstfeld, M., Taylor. A. M., 1997. Nonlinear aspects of goods-market arbitrage and adjustment: heckscher´s commodity points revisited. *J. Japanese Int. Economies* **11**, 441-479.

Peltzman, S., 2000. Price rise faster than they fall. *J. Polit. Economy* **108**, 466-502.

Piggott, N. E., Marsh, T. L., 2004. Does food safety information impact U.S. meat demand?. *Amer. J. Agr. Econ.* **86**, 154-174.

Sanjuán, A. I., Dawson, P. L., 2003. Price transmission, BSE and structural breaks in the UK meat sector. *Europ. Rev. Agr. Econ.* **30**, 155-172.

Serra, T., Goodwin, B. K., 2003. Price transmission and asymmetric adjustment in the Spanish dairy sector. *Appl. Econ.* **35**, 1889-1899.

Serra, T., Goodwin, B. K., 2004. Regional integration of nineteenth century U.S. egg markets. *J. Agr. Econ.* **55**, 59-74.

Serra, T., Gil, J. M., Goodwin, B. K., 2006. Local polynomial fitting and spatial price relationship: price transmission in EU pork markets. *Europ. Rev. Agr. Econ.* **33**,415-436.

Smith, M. E., Raavensway, E. O., Thompson, S. R., 1988. Sales loss determination in food contamination incidents: an application to milk bans in Hawaii. *Amer. J. Agr. Econ.* **70**, 513-221.

Tong, H., 1978. *On a threshold model in pattern recognition and signal processing*, (Ed) C. Chen. Sijhoff and Noonhoff, Amsterdam.

Tsay, R., 1989. Testing and modeling threshold autoregressive processes. *J. Amer. Statistical Assoc.* **84**, 231-240.

Von Cramon-Taubadel, S., 1998. Estimating asymmetric price transmission with the error correction representation: An application to the German pork market. *Europ. Rev. Agr. Econ.* **25**, 1-18.

Wang,Y., Liu, Q., 2006.Comparison of Akaike information criterion (AIC) and Bayesian information criterion (BIC) in selection of stock–recruitment relationships. *Fish. Res.* **77**, 220-225.

Ward, R. W., 1982. Asymmetry in retail, wholesale and shipping point pricing for fresh vegetables. *Amer. J. Agr. Econ.* **62**, 205-212.

Weinberger, M. G., Dillon, W. R., 1980. The effects of unfavorable product information. *Advances in consumer Res.* **7**, 528-532.

Wolffram, R., 1971. Positive measures of aggregate supply elasticities: some new approaches-some critical notes. *Amer. J. Agr. Econ.* **53**, 356-390.

World Organization for Animal Health, 2009. World animal health situation dataset. Accessed March 2009, available at http://www.oie.int/eng/info/en_esbmonde.htm.

Figure 1

Evolution of the number of BSE cases in Spain (2000-2007)

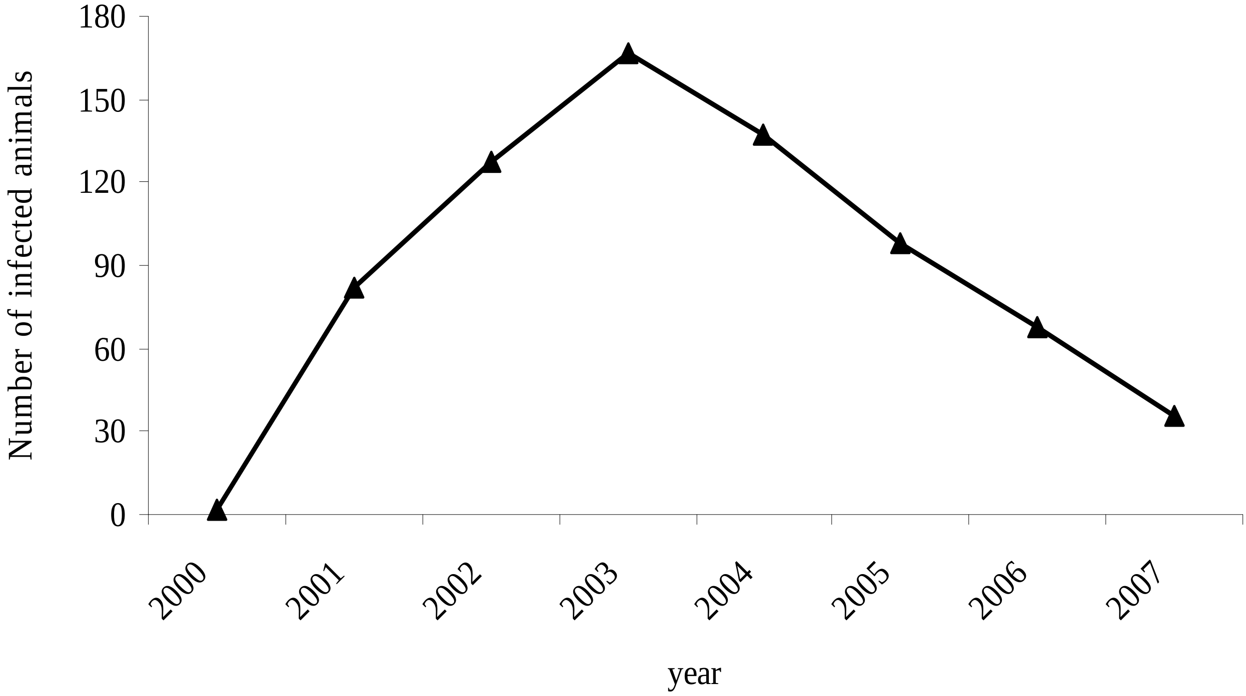

Figure 2

Food scare Information index.

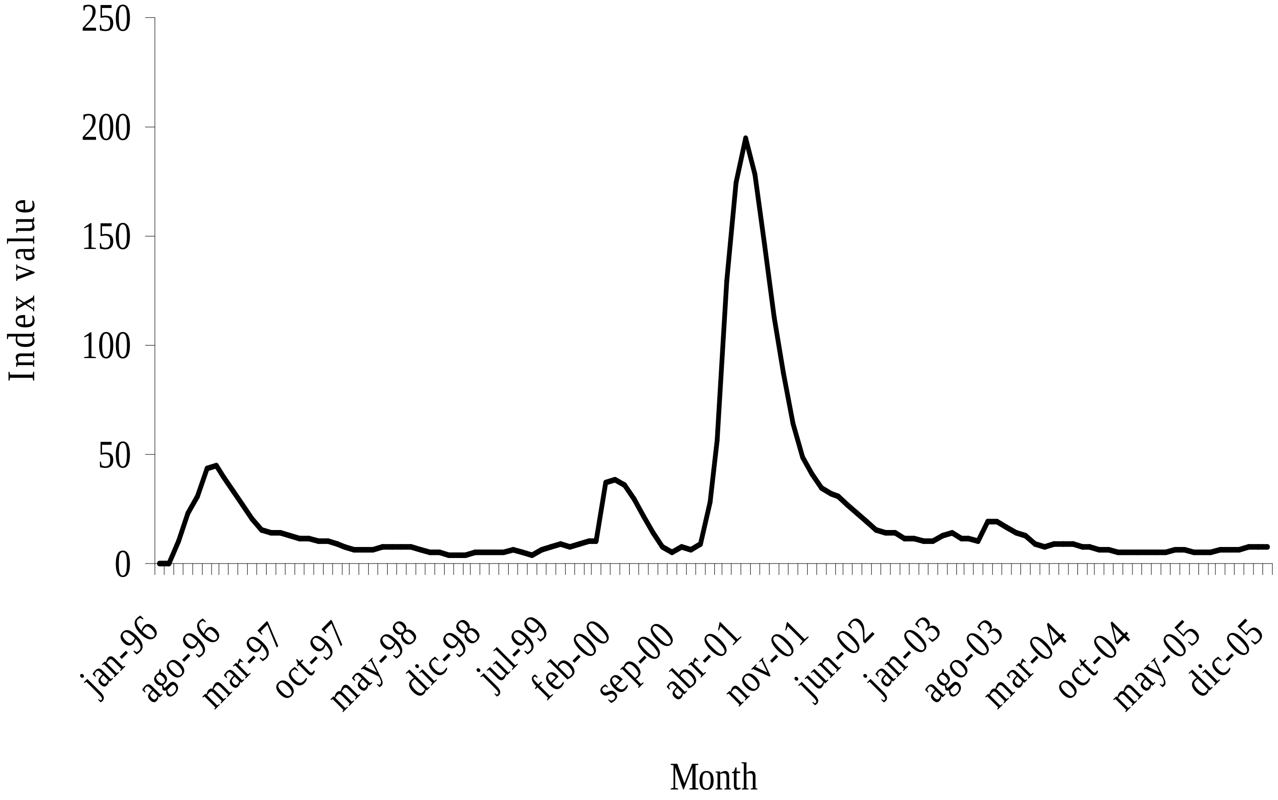

Figure 3

Timing of jumps among regimes

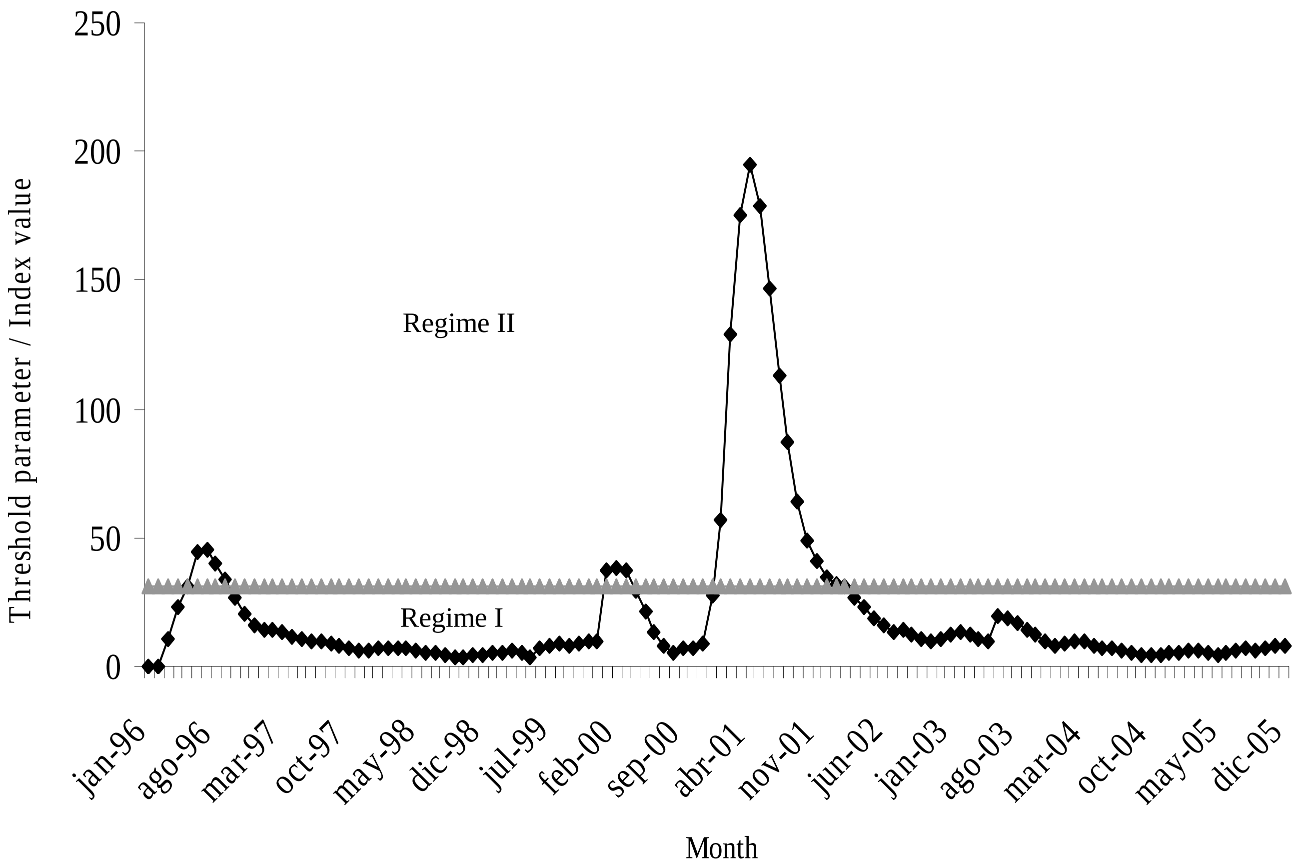

Table 1

OLS Estimates of the cointegrating relationship

Variable	Estimate	Standard Error
Intercept	-0.424**	0.190
Consumer Price	0.476**	0.095

Note: ** denotes statistical significance at the 5 per cent significance level

Table 2

Engle and Granger test for cointegration

	Test statistic (lag)	**10% critical value**
DF test	-3.256 (0)	-3.030

Notes: critical values are derived from Engle and Yoo (1987)

Table 3

RSVECM: Thresholds and Sup-LR test

Threshold (C)	Sup-LR test (p-value)
31.218	20.693 (0.048)

Table 4

RSVECM: parameter estimates

Dependent variables	**Producer price equation**		**Consumer price equation**	
	Regime I	Regime II	Regime I	Regime II
ΔPR_{t-1}	0.332**	0.132	0.769	1.177*
	(0.126)[a]	(0.130)	(0.653)	(0.652)
ΔPR_{t-2}	0.130	0.324**	-1.011	0.684
	(0.138)	(0.134)	(0.694)	(0.674)
ΔPR_{t-3}	0.175	0.322**	0.745	1.324*
	(0.132)	(0.144)	(0.667)	(0.724)
ΔCO_{t-1}	-0.004	-0.080**	-0.548**	-0.838**
	(0.027)	(0.033)	(0.135)	(0.166)
ΔCO_{t-2}	0.011	-0.085**	-0.162	-0.579**
	(0.029)	(0.043)	(0.144)	(0.217)
ΔCO_{t-3}	-0.016	-0.025	-0.188*	0.024
	(0.022)	(0.054)	(0.110)	(0.270)
ECT_{t-1}	-0.365**	-0.697**	0.242	-1.088
	(0.121)	(0.154)	(0.626)	(0.774)
Dummy	-0.111**		0.118	
	(0.036)		(0.182)	
Number of observations	Obs. in Regime I [100]		Obs. In Regime II [20]	

Notes: [a] Number in parentheses are standard errors.

*(**) denote statistical significance at the 10 (5) per cent level.