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Abstract. *In this paper we show how regime-switching vector error correction models can be used to assess the effects of food scares on price transmission along the food marketing chain. Our empirical implementation focuses on the impacts of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. Monthly prices are used over the 1996-2005 period. A BSE food scare index is developed in order to measure the scale of the food scare crisis. Results suggest that the crisis affects beef producers and retailers differently. While consumer prices are not found to adjust to the crisis, producer prices do, though only when the scale of the food scare reaches certain minimum levels.*

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

1. Introduction

The disease Bovine Spongiform Encephalopathy (BSE) generated a food scare when suspicions arose that it might be linked to new-variant Creutzfeld-Jakob Disease (vCJD). 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, thus increasing marketing margins in favor of the latter [1] 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) regarding BSE cases registered since 1987, after the United Kingdom, Ireland, Portugal and France [2].[1]

Second, the Spanish beef sector has economic and social relevance both within Spain and within the EU. Beef is one of the most important activities in the Spanish agricultural sector. It is the second most relevant meat sector, representing almost a fifth of the Spanish gross animal production [3]. Also, Spain occupies the fifth position in the EU rankings for bovine production with 671 thousand tons, representing 9% of total output [4]. 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 literature review of previous research. The third section is devoted to discussing econometric methods. The empirical application is presented in the fourth section. The discussion of the results is presented in the fifth section. The article concludes with the concluding remarks section.

2. 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 of the overall operation of the market [5]. 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 power 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 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.

[1] EU refers to EU-15.

Asymmetries in price transmission, which have received considerable attention in the literature, usually involve that positive shocks at one level of the market will elicit different responses at other levels compared to negative shocks (see, for example, [6], [7], [8]). A 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 market chain. A number of institutional and theoretical explanations have been suggested to explain asymmetries and these are reviewed in [9].

Though a great number of studies and reports about the BSE crisis have been made available recently, the literature on 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, [10], [11], [12], [13]). A few have tried to assess whether the BSE crisis altered the efficiency with which the sector operates [14]. 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.* [15] [1] 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.

Livanis and Moss [16] study the effects of food scares on price transmission in the U.S beef sector. Their methodological approach is similar to Lloyd *et al.* [15] 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. Jaenicke and Reiter [17] 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 among farm and retail markets. 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 demonstrate a threshold-type behavior and regime-switching is likely to be driven by the magnitude of the food scare. Markets may respond differently to distinct levels of a food scare, which could involve a change in marketing margins 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 and patterns of transmission 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. Our analysis is the first to focus on the impacts of BSE on price transmission within the Spanish beef market.

3. Econometric methods

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 [18] and later refined by Houck [19] and Ward [20]. These specifications have been criticized for ignoring the time-series properties of the (usually nonstationary) data. Von-Cramon Taubadel [21] extended the Wolffram specification to allow for an error correction term. More recently, Goodwin and Holt [5] proposed using threshold vector error correction models (TVECM) to allow error correction models to adequately capture nonlinear and threshold-type price adjustments.

Tong [22] originally introduced nonlinear threshold time series models. Tsay [23] developed a methodology to test for threshold nonlinearity based on autoregressions and to model threshold autoregressive processes. Balke and Fomby [24] combined error correction models and the threshold autoregressive model developed by Tsay [23] into a threshold error correction framework. These authors suggest implementing a grid search procedure to select threshold parameters that delineate different regimes that aim to minimize a sum of squared errors criteria.

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 [24] this analysis can be extended to a regime-switching autoregressive (RSAR) process. A three-regime RSAR can be expressed as:

$$\Delta v_t = \begin{cases} \gamma^{(1)} v_{t-1} & if \quad s_{t-d} \le c_1 \\ \gamma^{(2)} v_{t-1} & if \quad c_1 < s_{t-d} \le c_2 \\ \gamma^{(3)} v_{t-1} & if \quad c_2 < s_{t-d} \end{cases} \tag{1}$$

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 $d=1$). Parameters c_1 and c_2 represent thresholds that delineate the different regimes and $\gamma^{(m)}$, $m=1,2,3$ are speed of adjustment parameters that measure the rate at which prices adjust to disequilibriums from the long run equilibrium relationship. The RSAR can be alternatively expressed as $\Delta v_t = \gamma^{(1)} v_{t-1} d_{1t}(c_1, c_2, d) + \gamma^{(2)} v_{t-1} d_{2t}(c_1, c_2, d) + \gamma^{(3)} v_{t-1} d_{3t}(c_1, c_2, d) + u_t$, where $d_{1t}(c_1, c_2, d) = 1(s_{t-d} \le c_1)$, $d_{2t}(c_1, c_2, d) = 1(c_1 < s_{t-d} \le c_2)$, $d_{3t}(c_1, c_2, d) = 1(c_2 < s_{t-d})$. When the threshold variable is a lagged residual of the error correction term, $s_{t-d} = v_{t-d}$, the model is known as threshold autoregressive (TAR) model.

A Regime Switching Vector Error Correction Model (RSVECM) is a multivariate version of the RSAR model. The RSVECM allows one to uncover potential nonlinearities and asymmetries in the adjustment of individual prices and provides more information about short-run price dynamics[2]. Lo and Zivot [28] have suggested that the multivariate RSVECM should have higher power than univariate RSAR models. RSVECM 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 three-regime RSVECM can be expressed as follows:

$$\Delta P_t = \begin{cases} \alpha^{(1)} + \alpha_p^{(1)} v_{t-1} + \sum_{i=1}^{l} \alpha_i^{(1)} \Delta P_{t-i} + \varepsilon_t^{(1)} & if \quad s_{t-d} \le c_1 \\ \alpha^{(2)} + \alpha_p^{(2)} v_{t-1} + \sum_{i=1}^{l} \alpha_i^{(2)} \Delta P_{t-i} + \varepsilon_t^{(2)} & if \quad c_1 < s_{t-d} \le c_2 \\ \alpha^{(3)} + \alpha_p^{(3)} v_{t-1} + \sum_{i=1}^{l} \alpha_i^{(3)} \Delta P_{t-i} + \varepsilon_t^{(3)} & if \quad c_2 < s_{t-d} \end{cases} \tag{2}$$

where $P_t = (P_{1t}\ P_{2t})$ is the vector of prices being analyzed, $\alpha^{(m)}, \alpha_i^{(m)}$, $m=1,2,3$ 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. The RSVECM can be compactly expressed as:

$$\Delta P_t = \begin{cases} \boldsymbol{\alpha}^{(1)'}\, \mathrm{x}_{t-1} + \varepsilon_t^{(1)} & if \quad s_{t-d} \le c_1 \\ \boldsymbol{\alpha}^{(2)'}\, \mathrm{x}_{t-1} + \varepsilon_t^{(2)} & if \quad c_1 < s_{t-d} \le c_2\ , \\ \boldsymbol{\alpha}^{(3)'}\, \mathrm{x}_{t-1} + \varepsilon_t^{(3)} & if \quad c_2 < s_{t-d} \end{cases} \tag{3}$$

where:

$$\mathrm{x}_{t-1} = \begin{bmatrix} 1 \\ v_{t-1} \\ \Delta P_{t-1} \\ \vdots \\ \Delta P_{t-l} \end{bmatrix},$$

[2] When $s_{t-d} = v_{t-d}$, the RSVECM is known as TVECM. TVECM were recently discussed in [5], [25], [6], [7], [26], [27], and [8] among others.

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. We next estimate a three-regime RSVECM where the food scare information index is used as the threshold variable. Finally, we utilize the sup-LR statistic developed by Hansen and Seo [29] to test for a linear VECM against the alternative of a RSVECM.

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 parameters c_1 and c_2.[3] The variable relevant to threshold behavior has been usually assumed to be the (lagged) error correction term and the thresholds have been usually searched over negative and positive values of this term (see, for example, [6], or [7]).[4]

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 analyses have provided evidence that the BSE crisis had an impact on vertical price transmission processes ([15], [16]). The degree to which price transmission is affected by a food scare crisis is likely to depend on the scale of such crisis. A food scare information index (FSII) based on the number of news published on the topic is built in order to have a measure of this scale.[5] This FSII is used as a threshold variable. Hence, we implicitly assume that different levels of food scare can lead to different price behaviors and that the food scare index leads to switching between different price behavior regimes.

The first threshold is searched over the values of the FSII that are below the median, while the second threshold is searched over the values of the threshold variable that are above the median. The search is restricted in order to ensure an adequate number of observations for estimating the parameters in each regime. Recently, Serra and Goodwin [27] have considered two general grid search approaches in the selection of thresholds which may not be equivalent. The first approach involves minimization of the sum of squared errors or, alternatively, the trace of the covariance matrix for the system's residual errors. This approach has been implemented by a number of empirical analyses (see, [24] or [6]). The second approach maximizes a likelihood function (see, for example, [31] or [32]). As Serra and Goodwin [27] explain, the kernel of the likelihood function involves the log of the 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 approach of minimizing the logarithm of determinant of variance-covariance matrix of the residuals (Σ), the vectors of parameters $\alpha^{(i)}, i = 1,2,3$ are estimated by iterated seemingly unrelated regressions (SUR) method giving:

$$S(c_1, c_2, d) = 1n \left| \hat{\Sigma}(c_1, c_2, d) \right|. \tag{4}$$

where $\hat{\Sigma}(c_1, c_2, d)$ is a multivariate SUR estimate of $\Sigma = \text{var}(\varepsilon_t)$, conditional on (c_1, c_2, d) and $d = 1$. The vector of the errors is represented by $\varepsilon_t' = (\varepsilon_t^{(1)} \quad \varepsilon_t^{(2)} \quad \varepsilon_t^{(3)})$. In the second step, the least squares estimates of c_1 and c_2 are obtained as:

$$(\hat{c}_1 \, \hat{c}_2, 1) = \arg \min_{c_1, c_2} S(c_1, c_2, 1). \tag{5}$$

Final parameter estimates are given by $\hat{\alpha}^{(i)} = \hat{\alpha}^{(i)}(\hat{c}_1 \, \hat{c}_2, 1)$ and the estimations of the residual covariance matrix by $\hat{\Sigma} = \hat{\Sigma}(\hat{c}_1 \, \hat{c}_2, 1)$.

Finally, we test for the significance of the differences in parameters across relative regimes. The sup-LR test developed by Hansen and Seo [29] is used to test for a linear VECM against the alternative of a RSVECM. The model under the null is $\Delta P_t = \alpha' x_{t-1} + \varepsilon_t$, while the model under the alternative can be expressed as

[3] Several analyses that are based on threshold models have treated regime-switching as exogenous (see, for example, [30]). We adopt a more general model that incorporates this issue as endogenous.

[4] Alternatively, the error correction term can be segmented not according to whether it is greater or less than zero, but rather according to whether it is greater or less than a value that may differ from zero (see, for example, [29]).

[5] Details on how the index is built are provided in the next section.

$\Delta P_t = \boldsymbol{\alpha}^{(1)'} x_{t-1} d_{1t}(c_1, c_2, d) + \boldsymbol{\alpha}^{(2)'} x_{t-1} d_{2t}(c_1, c_2, d) \quad + \boldsymbol{\alpha}^{(3)'} x_{t-1} d_{3t}(c_1, c_2, d) + \varepsilon_t$. The sup-LR statistic can be computed in the following way:

$$LR = T\left(\ln\left|\hat{\Sigma}\right| - \ln\left|\hat{\Sigma}\right|(\hat{c}_1, \hat{c}_2, d)\right), \tag{6}$$

where $\hat{\Sigma}$ is the variance-covariance matrix of the residuals for the VECM, $\hat{\Sigma}(\hat{c}_1 \hat{c}_2, d)$ 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 bootstrap technique developed by Hansen and Seo [29].

More specifically, we carry out a parametric residual bootstrap method that requires the specification 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 the 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 at each period and derive the vector of series by recursion. The sup-LR test is computed for each replicated sample. The p-value can be determined as the proportion of simulations under the null for which the simulated LR statistic exceeds the observed statistic.

4. 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. These prices include 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 Spanish Ministry of Agriculture.[6]

As noted, health risks have received increasing attention among consumers in developed countries and created a strong relationship between food scares and 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, we construct a FSII that captures opinions concerning the degree of the consumer food scare and health risks. To do so, we build a FSII using a monthly count of newspaper articles on the BSE crisis appearing in a major Spanish newspaper, EL PAÍS.

Weinberger and Dillon [33] suggest that supporting articles (unfavourable news) may be more influential than a similar amount of non-supporting articles (favourable news). It has been also reported that a similar quantity of unfavourable news weighs far more heavily on consumer decision-making than favourable news (see, for example, [34], or [35]). However, Mazzochi [12] argues that discrimination between favourable and unfavourable articles is highly 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 [36] and Chern and Zuo [37], we do not weigh articles depending on whether they are favourable or unfavourable.

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.[7] The keywords searched were "Vaca(s) Loca(s)", "Encefalopatía Espongiforme Bovina", 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 of 20.8 news articles per month with a standard deviation of 44.12.

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 [37] to construct our monthly food scare information index, which assumes that published articles should have a finite duration and decaying effect as a source of consumer information. Specifically, Chern and Zuo [37] have extended the cumulative method used by Brown and Schrader [38] by building a new fat and cholesterol information index using a cubic weight function[8]. The monthly food scare information index used in our analysis is presented in figure 1.

[6] 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.

[7] 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.

[8] Different peak times and lifespans were considered, but results did not significantly differ. The final specification uses m=1 as peak time and considers a lifespan of 6 months.

5. Results

As noted above, our empirical analysis utilizes monthly farm-gate and consumer beef prices. It is also based upon a food scare index that is a measure of the degree of food scare caused by the BSE crisis on a monthly basis.

The empirical analysis is based upon logarithmic transformations of prices. Standard Dickey and Fuller [39] and KPSS [40] tests for each price series, provide evidence that all price series are integrated of order one I(1).[9] 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 (see, [41] p. 88, [42]).

After testing for unit roots, the Engle and Granger [43] test for cointegration is performed. 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 [6], [45], [46], or [16]). This is confirmed later in our analysis.

The Engle and Granger [43] test indicates the existence of a cointegration relationship among producer and consumer price series (see table 2). Both the AIC and SBC criteria recommend specifying the autoregressive process without lags. Other analyses have also found evidence of cointegration between producer and consumer prices for beef (see, for example, [15], [17], or [16]).

A RSVECM is then estimated by using sequential multivariate least squares in two steps[10]. The thresholds derived from the two-dimensional grid search and the sup-LR statistic are presented in table 3.[11] The sup-LR test statistic indicates that threshold effects are statistically significant. 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 three regimes, one corresponding to very low food scare index values (FSII<=6.323), another for intermediate values of this index (6.323 <FSII<=35.680), and a third regime corresponding to high levels of food scare (FSII>35.680). In table 4 we present the RSVECM parameter estimates across the different regimes.

Parameter estimates suggest that producer prices do all the adjustment to deviations from the long-run equilibrium relationship. Conversely, consumer prices are exogenous and do not adjust to system disequilibriums. The producer price adjustment, however, only occurs when the food scare information index takes high or intermediate values, while no adjustment takes place for low levels of food scare. In other words, when the weighted number of news articles published on the topic of the BSE is below 6.3 (regime I) producer prices do not adjust to shocks to the price system. However, when the weighted number of news is equal or greater than 6.3, producer prices respond to deviations from the long-run equilibrium (-42%). The response is even quicker (-51%) when the food scare information index reaches values greater than 35.7 (regime III). In summary, the BSE crisis seems to affect Spanish beef farmers and retailers differently, since all adjustments to deviations from equilibrium include the producer price, and only occur when the food scare information index takes intermediate or high levels.

Figure 2, presents the timing of jumps between the three alternative regimes.[12] The figure illustrates that regime I (low food scare) takes place during 1998 and 1999, as no case of BSE was reported during this period of time. The system returns to this regime in 2004, once the market reaches stability after the strong crisis that shook the system as a result of the discovery of the first BSE cases in Spain in 2000. A majority of observations fall into regime II (intermediate food scare) which generally 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 III (high food scare), 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. Since price adjustment is costly, results indicate that it only takes place after the food scare crisis has reached a minimum size (more than 6 weighted news articles on the topic published within a month). It is at this point when producer prices start to move back to the

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9 Results are available from the authors upon request.

10 The optimum number of lags was selected by considering the Breusch-Godfrey Lagrange Multiplier test for autocorrelation and the SBC and AIC criterion. Though the SBC and AIC criterion recommended the use of just one lag, three lags were 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 limited the number of simulations to 250 and we simplified the bootstrapped model to the consideration of one lag.

12 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 to forecast that the use of the error correction term instead of the index as the threshold variable may not have substantially altered the results.

equilibrium with a path that will be accelerated when more than 35 weighted news articles per month appear in the press, i.e., when consumers read about the crisis at least once a day on average.

6. Concluding remarks

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 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 as a threshold variable.

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 yet.

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 [37] 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 disequilibriums caused by the food scare. Conversely, producer prices are endogenous and adjust as a response to the crisis. This adjustment only occurs when the food scare information index takes high or intermediate levels, while no adjustment takes place with low levels of food scares. 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 (see, for example, [16], [15], [1] and [17]). Finally, the sup-LR test statistic indicates that threshold effects are statistically significant.

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.* in [47] explained that using non-parametric techniques can overcome the limitations involved by parametric threshold models and thus it would be also useful to replicate our analysis using these non-parametric techniques. This would allow determining to what extent our results are subject to the specific functional forms used in the analysis.

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Figure 1 Food scare Information index based on Chern and Zue method.

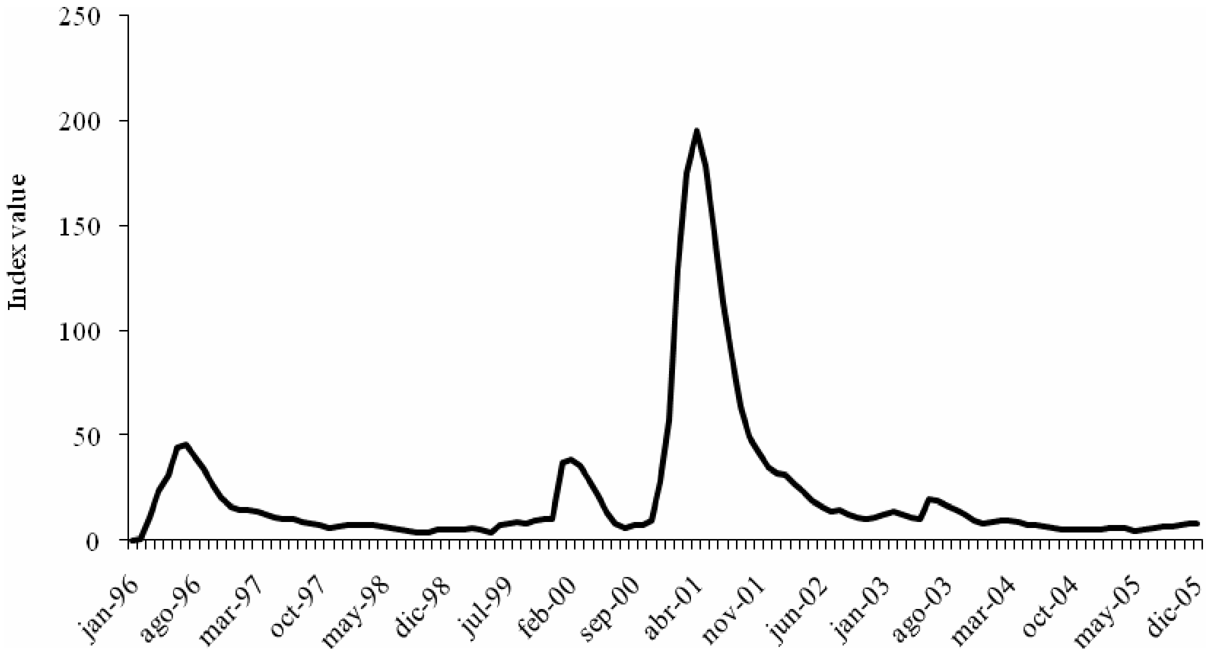

Figure 2 Timing of jumps among the regimes

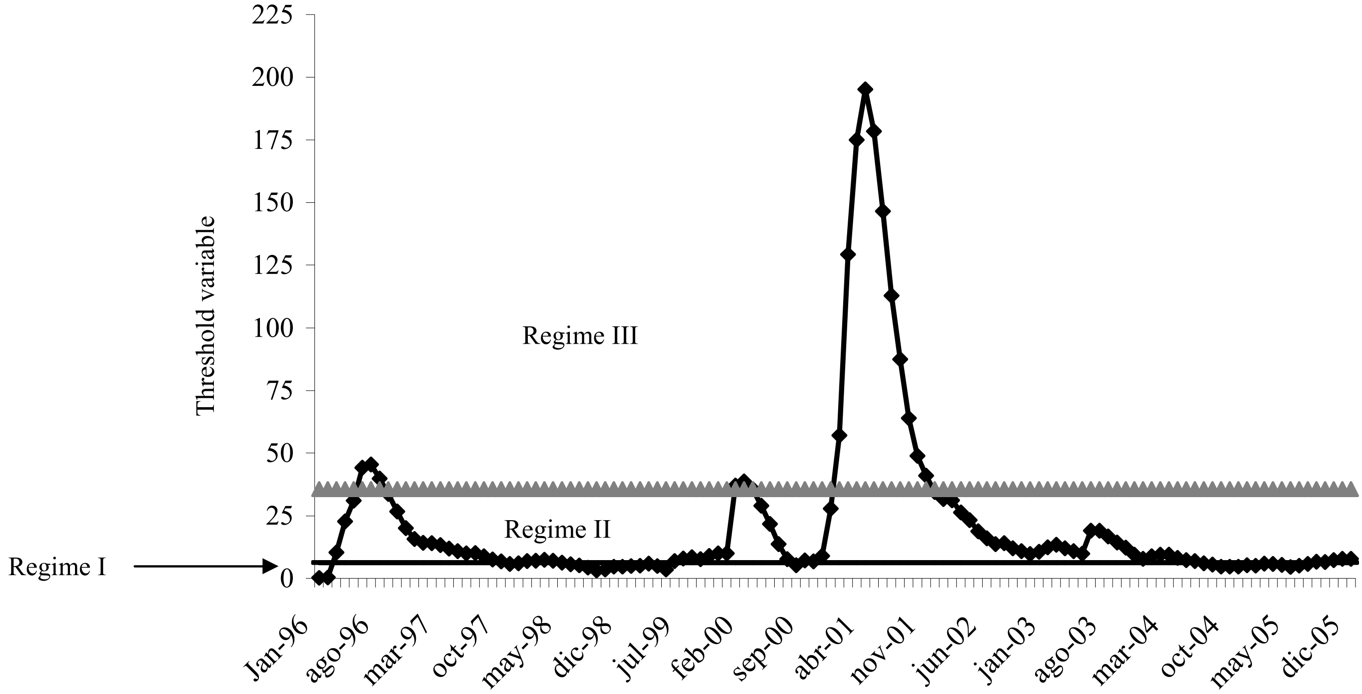

Table 1 OLS Estimates of the Cointegrating Relationship (normalized by the producer price)

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

	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

C1: first threshold	C2: second threshold	Sup-LR test (p-value)
6.323	35.680	36.572 (0.0240)

Table 4 RSVECM: parameter estimates

Dependent variables	Producer price equation			Consumer price equation		
	Regime I	Regime II	Regime III	Regime I	Regime II	Regime III
ΔPR_{t-1}	0.289	0.348**	0.074	0.542	0.785	0.626
	(0.281)	(0.139)	(0.155)	(1.325)	(0.683)	(0.732)
ΔPR_{t-2}	-.034	0.214	0.472**	-2.037	-0.154	-0.443
	(0.245)	(0.163)	(0.171)	(1.153)	(0.771)	(0.806)
ΔPR_{t-3}	0.263	0.088	0.547**	-0.264	0.643	0.433
	(0.293)	(0.154)	(0.183)	(1.380)	(0.735)	(0.863)
ΔCO_{t-1}	-0.002	-0.020	-0.055	-0.870**	-0.612**	-0.642**
	(0.069)	(0.030)	(0.040)	(0.329)	(0.141)	(0.192)
ΔCO_{t-2}	-0.032	0.005	-0.138	-0.552	-0.376**	0.062
	(0.077)	(0.028)	(0.076)	(0.365)	(0.132)	(0.361)
ΔCO_{t-3}	-0.025	-.0174	-0.020	-0.453	-0.249**	0.239
	(0.065)	(0.024)	(0.060)	(0.306)	(0.116)	(0.282)
ECT_{t-1}	-0.303	-0.418**	-0.510**	-1.437	1.413	-0.829
	(0.226)	(0.151)	(0.157)	(1.068)	(0.741)	(0.741)
Number of observations	Obs. in Regime I [31]		Obs. In Regime II [72]		Obs. in Regime III [17]	

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

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