ASYMETRIC PRICE TRANSMISSION IN THE SPANISH LAMB SECTOR

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Paper prepared for presentation at the XI[th] Congress of the EAAE
(European Association of Agricultural Economists),
'The Future of Rural Europe in the Global Agri-Food System'
Copenhagen, Denmark 24-27 August 2005

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Abstract

This paper aims to investigate the non-linear adjustments of prices between farm and retail prices in the lamb sector in Spain. The methodology used is based on the multivariate approach to specify and estimate a three-regime Threshold Autoregressive Model. Results indicate that in the long-run price transmission is perfect and any supply or demand shocks are fully transmitted along the marketing chain. In the short-run, price adjustments between the farm and the retail levels are asymmetric and are representative of a demand-pull transmission mechanism. On the other hand, retailers benefit from any shock, whether positive or negative, that affects supply or demand conditions.

Key words: Asymmetries, lamb, Spain, price transmission

1. Introduction

The issue of price transmission in a vertical sector has been the subject of much research. Particularly, the asymmetric price adjustments have been one of the main research interests among agricultural economists. While asymmetric price responses do not necessarily emanate from market inefficiencies, evidence of symmetric, rapid price responses is clear evidence of an efficient market. It is a common feeling that retail prices do not react very quickly to changes in market conditions. A good example could be the situations where retail prices remain sticky in spite of decreases of input prices due to primary production increases (Borenstein et al., 1997; Peltzman, 2000). In this situation the retail price will not be equal to the marketing clearing price, generating excess supply and consumers will not benefit for declining farm prices suggesting a redistribution of consumer welfare.

As the standard theory of markets has the general implication of symmetric price adjustment, the knowledge regarding price asymmetry has come mainly from the accumulation of empirical results. Different potential explanations have been given for the existence of asymmetries in price adjustments: the market power at the retail level[1], adjustment costs at the retail level, input substitution at the processing level, stocks at both the production and retail level, production lags at the processing level and public intervention. However, only in a few cases (Peltzman, 2000) there has been an attempt to link asymmetries with any of the mentioned potential explanations. In any case, before explanations can be given for specific markets, the first step is to analyse the existence of such asymmetric price adjustments.

The standard approach to test for asymmetries in price transmission relies on variations of a model first developed by Wolffram (1971) and later modified by Houck (1977) and Ward (1982). Recent developments in time series analyses have modified the methodological framework to tackle with this issue. Von Cramon-Taubadel (1998) showed that the traditional approach was inconsistent

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[1] Although asymmetries have been linked to non-competitive behaviour, this is not necessarily true. McCorriston et al. (2001), with formal grounding in rational firm conduct, showed that in presence of market power price changes could be greater or less than the competitive benchmark case depending on the interaction between such market power and returns to scale. If the industry cost function is characterised by decreasing returns to scale the damping effect of market power is reinforced. On the other hand, if the cost function is characterised by increasing returns to scale the market power effect is offset.

with cointegration among prices, specifying a linear Error Correction Model. However, the existence of transaction costs may generate non linear price adjustments, making threshold models more suitable to analyse asymmetries in price transmission (Azzam, 1999).

In this paper we investigate, using time series data, the existing asymmetries in the price transmission mechanism between farm and retail marketing channels in the Spanish lamb markets. Particularly, we will focus our study to answering the question of whether Spanish lamb farmers benefit or not from unanticipated positive and negative supply or demand shocks. By using time series observations we will be able to study the inevitable dynamic aspects of price transmission along the Spanish lamb sector. A Threshold Vector Error-Correction Model (TVECM) will be specified and non-linear impulse response functions will be calculated to tackle with this issue. Finally, results will be discussed taking into account the specific characteristics of the Spanish lamb sector.

Spain is the second largest lamb producer within the European Union (EU) just behind the United Kingdom. It represents around 5% of the Spanish Final Agricultural production and 11% of Final Livestock Production. Although the Common Market Organisation has a set of operating rules which may influence the movement of sheep on to or off a holding they do not place any physical constraints on producers adapting production to meet consumer needs, etc. However, the way in which the calculation of the premium is made has the potential to dissuade those flocks which produce less lamb per ewe then the average from responding to market signals as they have the real potential to achieve a lower income, even at higher prices per lamb, because of the potential for the premium payment to decline as market prices improve. Although producers are unlikely to deliberately sell product at low prices they will not make a significant effort to change their system to capitalise on higher market prices for different qualities of product or at different times of year. Conversely however, because the premium is equal for all ewes, individual producers who achieve better than average market prices through improved quality or other market initiatives and who produce more lamb per ewe than the standard will potentially achieve higher margins. Consequently farmer prices are not primarily determined by the ewe premium and producers have incentives to adapt to changing market conditions.

To achieve the above-mentioned objective, the rest of the paper is organised as follows. Section 2 provides a description of the methodological approach used in the paper. Section 3 reports our empirical results. Finally, section 4 closes the paper with some concluding remarks.

2. Modeling nonlinear adjustments

Several studies attempting to measure asymmetric price transmission focused of the estimation of the following model (which constitutes a variation of the econometric specification introduced by Wolffram (1971) and redefined by Houck (1977) and Ward (1982)):

$$\Delta p_{1t} = \alpha_0 + \sum_{i=1}^{k} \gamma_{1i} I_t \Delta p_{2t-i} + \sum_{i=1}^{k} \gamma_{2i} (1 - I_t) \Delta p_{2t-i} + \varepsilon_t$$

where Δp_{1t} and Δp_{2t} represent changes in retail and wholesale prices, respectively, and I_t is an indicator function that is equal to one if $\Delta p_{2t-1} \leq 0$ and zero in other case. This model allows us to test if the response of retail prices differs depending on whether wholesale prices increase or decrease (see, for example, Hahn (1990), Kinnucan and Forker (1987) and Bailey and Brorsen (1989)).

However, results from the empirical models used by the above authors to investigate asymmetries in price transmission have been criticised for the following reasons: i) this specification assumes that the causality goes from wholesale to retail prices only; ii) this model has been used without adequately analysing the time series properties of data. Price levels often exhibit a non-stationary covariance property which, as a consequence, may bias causality tests and lead to autocorrelation problems in the asymmetric price response function (Boyd and Brorsen, 1988, and Kinnucan an Forker, 1987). On the other hand, if the price series are cointegrated, the specification of a model in first differences is biased as a result of the misspecification of the long-run relationships between prices. Von Cramon-Taubadel (1998) showed that the traditional econometric specification used to test for asymmetric price transmission is inconsistent with cointegration. He proposed an alternative specification of the Wolffram-Houck model based on the error correction representation, and taking into account the procedure approach suggested by Granger and Lee (1989). Balke et al. (1998) and Frost and Bowden (1999) also use an error correction model to test for asymmetric adjustment. However, these applications are based on linear error correction models. The presence of fixed costs of adjustment along the food chain may generate non-linear reactions; that is to say, price adjustments may be different depending both on the magnitude and the sign of the initial shock. In other words, it is not unrealistic to suppose that only when the initial shock surpasses the critical threshold do economic agents react to it. Balke and Fomby (1997) present a model that allows for non-linear adjustment to the long run equilibrium by introducing the concept of *threshold cointegration*.

In this context, two different methodological approaches have been developed. The first one is based on an univariate version of the bivariate threshold cointegration models described by Balke and Fomby (1997), Enders and Granger (1998) and Enders and Siklos (2001). In a similar way to the two-step Engle and Granger cointegration approach, this univariate procedure analyses the threshold behaviour of the univariate cointegrating residual implied by the prices spread, equal to log price difference. It assumes that one of the two prices is exogenous and that only the adjustments to the equilibrium change with regimes, while the autoregressive parameters of the model remain constant.

The second approach to test for threshold cointegration has been suggested by Hansen and Seo (2002) and Lo and Zivot (2001). As in Balky-Fomby (1997), the analysis of threshold behaviour is based on a bivariate Vector Error Correction Model (VECM) with one cointegrating vector. Hansen and Seo (2002) and Lo and Zivot (2001) indicate that the analysis of threshold behaviour in the bivariate model allows us to uncover potential nonlinearities and asymmetries in the adjustment of individual prices and provide more information regarding the dynamic of the data. In addition, as such a procedure utilises the full structure of the model, it should have higher power, provided the model is true, than univariate procedures, which ignore the restrictions imposed by the multivariate structure. This is the approach followed in this paper, which is described in the following section.

2.1 Threshold cointegration

Studies on price transmission using threshold error correction models (either univariate or bivariate) have either considered one threshold (λ^1) to separate the adjustment process into two regimes (Balke and Fomby, 1997; Enders and Granger, 1998, Abdulai, 2000 and 2002, Hansen and Seo, 2002) or two thresholds (λ^1 and λ^2) to separate the adjustment process into three regimes (Obstfeld and Taylor, 1997; Goodwin and Piggott, 2001; Serra and Goodwin, 2002, Meyer, 2003, etc.). Several authors suggest that a price adjustment model with three regimes separated by two thresholds has more economic sense than a two regime model with only one threshold (Meyer, 2003).

In this paper we start our analysis by considering a general three-regime Threshold Vector Error Correction Model ($TVECM_3$) to analyse price dynamics along the Spanish lamb chain. Let $P_t=(P_{1t},P_{2t})'$ be the log price of a good at two different levels of the marketing channel, assuming that P_t is a vector of I(1) time series which is cointegrated with a common cointegrating vector $\beta' = (1,-\beta_2)$. The linear VECM representation of order k of P_t can be written as:

$$\Delta P_t = \alpha[\omega_{t-1}(\beta)] + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \varepsilon_t \qquad (1)$$

where $\omega_t(\beta) = \beta' P_{t-1}$ is the cointegrating vector evaluated at the generic value $\beta=(1,-\beta_2)'$; Γ_i, i= 1, 2… are (2×2) matrices of short-run parameters; α is a (2×2) matrix; and ε_t is a vector of error terms that are assumed to be independently and identically Gaussian distributed, with a covariance matrix Σ which is assumed to be positive definite. β is the cointegrating vector which is commonly interpreted as the long-run equilibrium relation between the two prices in P_t , while α gives the weights of the cointegration relationship in the VECM equations.

Following Lo and Zivot (2001), a three-regime threshold Vector Error Correction Model ($TVECM_3$), can be written as:

$$\Delta P_t = \begin{cases} \alpha^1 \omega_{t-1}(\beta) + \sum_{i=1}^{k-1} \Gamma_i^1 \Delta P_{t-i} + \varepsilon_t^1, & \text{if} \quad \omega_{t-1}(\beta) < \lambda^1 \\ \alpha^2 \omega_{t-1}(\beta) + \sum_{i=1}^{k-1} \Gamma_i^2 \Delta P_{t-i} + \varepsilon_t^2, & \text{if} \quad \lambda^1 \le \omega_{t-1}(\beta) \le \lambda^2 \\ \alpha^3 \omega_{t-1}(\beta) + \sum_{i=1}^{k-1} \Gamma_i^3 \Delta P_{t-i} + \varepsilon_t^3, & \text{if} \quad \omega_{t-1}(\beta) > \lambda^2 \end{cases} \tag{2}$$

where $\omega_t(\beta)$ is the threshold variable which represents the residual of the equilibrium relationship (i.e. a deviation from equilibrium), and $\lambda = \begin{pmatrix} \lambda^1 & \lambda^2 \end{pmatrix}$ are the threshold parameters that delineate the different regimes.

As can be observed, the $TVECM_3$ in (2) specifies that the adjustment towards the long-run equilibrium relationship is regime specific. This model says that the dynamic adjustment of P_{it} depends on the magnitude of $\omega_t(\beta)$. A special case of the TVECM given in (2) occurs if price changes are smaller than transaction costs. In this case, prices will not adjust in the second regime (in the middle one) implying that prices are not cointegreted, that is, $\alpha^2=0$. The resulting model is the so-called Band-TVECM. In this case, if $\omega_t(\beta)$ is within the band then prices are not cointegrated and P_t follows a VAR(k) without a drift. However, in the outer bands economic forces push prices moving together implying cointegration with different adjustment coefficients. If $\omega_t(\beta) > \lambda^2$ ($\omega_t(\beta) < \lambda^1$), then the cointegrating vector reverts to the regime- specific mean with adjustment coefficient[2] ρ^3 (ρ^1) while ΔP_t adjusts to the long run equilibrium with a speed of adjustment vector α^3 (α^1). It is important to emphasise that the speed of adjustment of prices in the outer bands can be different for each element of P_t. The resulting model is given by:

$$\Delta P_t = \begin{pmatrix} \Delta P_{1t} \\ \Delta P_{2t} \end{pmatrix} = \begin{cases} \begin{pmatrix} \alpha_1^1 \\ \alpha_2^1 \end{pmatrix} \omega_{t-1}(\beta) + \sum_{i=1}^{k-1} \Gamma_i^1 \Delta P_{t-i} + \varepsilon_t^1, & \text{if} \quad \omega_{t-1}(\beta) < \lambda^1 \\ \sum_{i=1}^{k-1} \Gamma_i^2 \Delta P_{t-i} + \varepsilon_t^2, & \text{if} \quad \lambda^1 \le \omega_{t-1}(\beta) \le \lambda^2 \\ \begin{pmatrix} \alpha_1^3 \\ \alpha_2^3 \end{pmatrix} \omega_{t-1}(\beta) + \sum_{i=1}^{k-1} \Gamma_i^3 \Delta P_{t-i} + \varepsilon_t^3, & \text{if} \quad \omega_{t-1}(\beta) > \lambda^2 \end{cases}$$

However, the above model does not say anything about the direction of causality and the asymmetric adjustment process. Information about such features is provided by the α_i^j coefficients. In general, we expect $\alpha_1^j \le 0$ and $\alpha_2^j \ge 0$, that is, prices adjust to the long-run equilibrium when price changes are large. In any case, and assuming a two-price system (i.e. prices at retailing (p_1) and producer levels (p_2)), other interesting cases in this context are the following:

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[2] The adjustment coefficient is obtained as follows: $\rho^j = 1 + \beta' \alpha^j = 1 + \begin{bmatrix} 1 & -\beta_2 \end{bmatrix} \begin{bmatrix} \alpha_1^j \\ \alpha_2^j \end{bmatrix} = 1 + \alpha_1^j - \beta_2 \alpha_2^j$

- $\alpha_1^j = 0$	Retail prices do not respond to changes in the marketing margin. Retail prices are sticky relative to producer prices
- $\alpha_i^1 = 0$ and $\alpha_i^3 \neq 0$ i=1,2	Prices respond to positive shocks, but negative shocks in the marketing margin are allowed to persist
- $\left|\alpha_i^1\right| > \left|\alpha_i^3\right|$ i = 1,2	The adjustment towards the long-run equilibrium relationship between producer and retail prices is faster when changes in deviations are negative (i.e. producer prices rise and the marketing margin decreases) than when they are positive (i.e. producer prices decline and the marketing margin increases)

The three-regime TVECM given in (2) can be compactly expressed as the following multivariate regression model:

$$\Delta P'_t = X'_{t-1}A^{(1)}I_t^1(\lambda) + X'_{t-1}A^{(2)}I_t^2(\lambda) + X'_{t-1}A^{(3)}I_t^3(\lambda) + \varepsilon_t \qquad (3)$$

where:

$\lambda = (\lambda^1 \quad \lambda^2)$

$I_t^j(\lambda) = I(\lambda^{(j-1)} < \omega(\beta) < \lambda^{(j)})$ is a heavyside indicator function such that I(A)=1 if A is true and 0, otherwise.

$$X'_{t-1} = \left(\omega_{t-1}(\beta) \quad \Delta P'_{t-1} \quad \cdots \quad \Delta P'_{t-k+1}\right) \qquad A^{(i)} = \begin{pmatrix} \alpha^i \\ \Gamma_1^i \\ \vdots \\ \Gamma_{k-1}^i \end{pmatrix} \text{is a } (k+1)\times 2 \text{ matrix}$$

Note that when threshold parameters (λ^1 and λ^2) are both fixed (known a priori), the model is linear in the remaining parameters. In such circumstances, and under the assumption that errors ε_t are iid gaussian, parameters in model (3) can be estimated by multivariate least squares.

However, in general, the threshold parameters (λ^i's) are unknown and need to be estimated along with the remaining parameters of the model. Lo and Zivot (2001) propose a strategy which combine the Hansen's (1999) approach to estimate two- and three-regime univariate TAR models and the Tsay 's (1998) procedure to estimate multivariate TVECM. This strategy consists of the following steps. In the first step, a two-dimensional grid searches are carried out to estimate the threshold parameters (λ^1, λ^2) under the following assumptions: i) threshold parameters are such that $\lambda^i \in \Gamma_2, i = 1,2$ where $\Gamma_2 = \left\{(\lambda^1, \lambda^2) : -\infty < \lambda^L < \lambda^1 < \lambda^2 < \lambda^U < \infty\right\}$ (this assumption restricts all threshold parameters to lie in the bounded subset $[\lambda^L, \lambda^U]$), and ii) the search is restricted to ensure an adequate number of observations for estimating the parameters in each regime.

In practice, the analysis is conducted by imposing an ad-hoc bound for the number of observations in each regime. Letting T_i the number of observations in regime i, Hansen (1999) suggests constraining the threshold parameters such that $T_i / T \geq \pi_0$, with typically (π_0) set to 0.1. Conditional on $\lambda = (\lambda^1 \quad \lambda^2)$ the TVECM (3) is linear in the $A^{(i)}$'s and may be estimated by sequential multivariate least squares minimising:

$$S_3(\lambda^1, \lambda^2) = \ln\left|\hat{\Sigma}(\lambda^1, \lambda^2)\right| = \ln\left|\frac{1}{T}\sum_{t=1}^{T}\hat{\varepsilon}_t(\lambda)\hat{\varepsilon}_t(\lambda)'\right| \qquad (4)$$

where $\hat{\Sigma}_3(\lambda^1, \lambda^2)$ is the estimated covariance matrix of model (3) conditional on (λ^1 and λ^2).

In the second step, the threshold parameters can be estimated through the following optimisation program[3]:

$$(\hat{\lambda}^1, \hat{\lambda}^2) = \arg\min_{\lambda \in [\lambda^L, \lambda^U]} \left(S_3(\lambda^1, \lambda^2)\right) \tag{5}$$

The final parameter estimates of the TVECM (2) can be computed as $\hat{A}^{(i)} = \hat{A}^{(i)}(\hat{\lambda}^1, \hat{\lambda}^2)$ and the residual covariance matrix is given by $\hat{\Sigma}_3(\hat{\lambda}) = \hat{\Sigma}_3(\hat{\lambda}^1, \hat{\lambda}^2)$. Tsay (1998) shows that the conditional least squares estimators of the TVECM are strongly consistent as the sample size increases ($\hat{A}^i \to A^i, \hat{\lambda}^i \to \lambda^i$, and $\hat{\Sigma}_3(\hat{\lambda}) \to \Sigma$) and that the parameters of $A^{(i)}$'s matrices are asymptotically normally distributed.

The third step consists of testing if the dynamic behaviour and the adjustment towards the long-run equilibrium relationship is linear or exhibits threshold non-linearity. Several univariate and multivariate test for linearity that have power against the threshold alternative have been proposed in the literature (Balke and Fomby (1997), Hansen (1997, 1999), Hansen and Seo (2002), Tsay (1998).

Lo and Zivot (2001) suggest the Hansen's method for testing linearity in univariate TAR models based on nested hypothesis tests and which can be easily extended to test linearity in multivariate TVECMs. They propose the sup-LR statistic:

$$LR_{13} = T\left(\ln\left|\hat{\Sigma}\right| - \ln\left|\hat{\Sigma}_3(\hat{\lambda})\right|\right) \tag{6}$$

where $\hat{\Sigma}$ and $\hat{\Sigma}_3(\hat{\lambda})$ are the residual covariance matrices of the VECM and three-regime TVECM, respectively.

The statistic to test such a hypothesis suffers from the problem of the so-called unidentified nuisance parameters under the null hypothesis. In other words, the non-linear model contains certain parameters which are not restricted under the null hypothesis and which are not present in the linear model. Consequently, the conventional statistical theory cannot be applied to obtain the asymptotic distribution of the statistics (see Davies, 1987; Hansen, 1999 and Hansen and Seo, 2002). Given that the test statistic has a non-standard distribution, Hansen (1999) and Hansen and Seo (2002) suggest using the fixed regressor bootstrap or, alternatively, a parametric residual bootstrap algorithm, to compute the p-value for the linearity tests.

Once the presence of threshold effects is confirmed, in the empirical analysis there are several questions that they would have to be answered before allowing the researcher to interpret results. In this context, the most important, with no doubt, is to determine which kind of threshold model is more appropriated for the data (number of regimes, TVECM or Band-TVECM, and symmetric or asymmetric threshold model). Two approaches have generally been considered to determine the appropriate threshold specification. The first approach uses a model selection criterion (AIC, SBC, etc.) to determine the best specification form the data (Tsay, 1998). Following Hansen (1999), Lo and Zivot (2001) consider nested hypothesis tests based on an unrestricted estimation of the TVECM. They consider, first, the determination of the number of regimes. Thus, in order to test the null of a $TVECM_2$ (two-regime model) against the alternative of a $TVECM_3$ (three-regime model) they propose the following Likelihood Ratio (LR) statistic:

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[3] The grid research minimizes the log determinant of the residual covariance matrix of the TVECM, which is analogous to maximizing a standard LR test.

$$LR_{2,3} = T\left(\ln\left|\hat{\Sigma}_2(\hat{\lambda})\right| - \ln\left|\hat{\Sigma}_3(\hat{\lambda})\right|\right) \quad (7)$$

where $\hat{\Sigma}_2(\hat{\lambda})$ and $\hat{\Sigma}_3(\hat{\lambda})$ are the estimated residual covariance matrices from the unrestricted two-regime $TVECM_2$ and three-regime $TVECM_3$, respectively. The asymptotic distributions of $LR_{2,3}$ are non-standard and bootstrap methods can be used to compute approximate p-values.

Once the number of regimes has been established, they propose specification tests for the Band-TVECM. Since the estimated threshold parameters from the TVECM are superconsistent, as mentioned previously, then a Wald test can be used, by defining appropriate restrictions on the TVECM parameters, which follows an asymptotic chi-square distribution.

2.2 Non-linear impulse response functions

Once the TVECM has been estimated, it is useful to analyse the short-run dynamic behaviour of the variables by computing the impulse response functions. This can be particularly suitable for studying the time path response of variables to unexpected shocks at time t. However, given that the non-linear time series model does not have a Wald representation, computing the IRF for these types of models is not an easy task. In addition, as discussed in Koop et al. (1996), the complications arise because in non-linear models: i) the effect of a shock depends on the history of the time series up to the point where the shock occurs; and ii) the effect of a shock depends on the sign and the size of the shock. As a consequence, in non-linear models impulse response functions depend on the combined magnitude of the history $P_{t-1}=\omega_{t-1}$ and the magnitude of the shock δ (relative to the threshold value λ)

The Generalised Impulse Response Functions (GIRF) introduced by Koop et al. (1996) and Potter (1995) offer a useful generalisation of the concept of impulse responses to non-linear models. Their analysis focused on the asymmetric response of the variables to one standard deviation of both positive and negative shocks. The Non-linear Impulse Response Functions (NIRF) are defined in a similar manner to traditional GIRF, except for replacing the standard linear predictor by a conditional expectation. Hence, the NIRF for a specific shock $\varepsilon_t = \delta$ and history $P_{t-1}=\varphi_{t-1}$ (the history of the system) is defined as:

$$\begin{aligned} NIRF(n,\delta,\varphi_{t-1}) = {} & E\left[P_{t+n} \mid \varepsilon_t = \delta, \varepsilon_{t+1} = \ldots = \varepsilon_{t+n} = 0, \varphi_{t-1}\right] \\ & - E\left[P_{t+n} \mid \varepsilon_t = 0, \varepsilon_{t+1} = \ldots = \varepsilon_{t+n} = 0, \varphi_{t-1}\right] \text{ for } n = 0,1,\ldots N \end{aligned} \quad (8)$$

Taking into account this definition, it is clear that the NIRF is a function of $\delta \in \varepsilon_t$ and $\varphi_{t-1} \in \Omega_{t-1}$ (Ω_{t-1} is the history or information set at t-1 used to forecast future values of Pt). Given that δ and φ_{t-1} are realisations of the random variables Ω_{t-1} and ε_t, Koop et al. (1996) stress that NIRF themselves are realisations of random variables given by:

$$NIRF(n,\varepsilon_t,\Omega_{t-1}) = E\left[P_{t+n} \mid \varepsilon_t, \Omega_{t-1}\right] - E\left[P_{t+n} \mid \Omega_{t-1}\right] \quad (9)$$

From (9), there are a number of alternative ways to calculate the NIRF, depending on the research objectives. For instance, in this study we have considered it relevant to assess the responses of wholesale (retail) prices to shocks in retail (wholesale) prices under different evolution price regimes, and under different sizes and signs of the initial shock. Thus, the NIRF can be used to evaluate the degree of asymmetric responses over time.

3. Empirical analysis

3.1. Data and preliminary analysis

In this section we perform the multivariate threshold cointegration approach described above to analyse the price transmission mechanism along the Spanish lamb marketing chain. Empirical specification TVECM involves the following steps: i) under the assumption of prices non-stationarity, the first step consists of testing for cointegration and estimating the cointegrating relationships; ii) if cointegration is found, the next step consists of determining whether the dynamics of the data can be described by threshold-type nonlinearities; iii) estimation and evaluation of the bivariate threshold error correction model (TVECM), and iv) non-linear Generalised Impulse Response functions are calculated in order to analyse the response of each prices to unanticipated positive and negative shocks. Each of these steps is addressed in turn in this section.

As mentioned in Section 3, our empirical analysis uses weekly data of farmer prices (FP), and retail (RP) prices along the period 1993-2002. All variables are expressed in natural logarithms. For cointegration analyses among prices, it is common to use logarithms because otherwise, with trending data, the relative error is declining through time (Banarjee et al., 1993). On the other hand, Tiffin and Dawson (2000) suggest that the logarithmic transformation is appropriate because the variance is related to the mean and the relative error is constant for the series in levels. From an economic point of view, this transformation allows to relate prices in terms of percentage variations instead of absolute changes.

Previous to the cointegration analysis among the price series, we first examine their stochastic time series properties. Seasonality has been investigated by implementing seasonal unit root tests for weekly data following the procedure suggested by Cáceres (1996) and Cáceres et al. (2001)[4]. Results from these statistics clearly suggest that seasonality is deterministic for the three price series. Accordingly, the systematic component of seasonality, to be parsimonious, has been adequately captured by using a Fourier-type series expansion.

Time series univariate properties have been examined by using unit root tests. As in small samples such tests have limited power, two alternative unit root tests developed by Elliot et al., (1996) and Ng and Perron (2001) as well as the stationary test from Kwiatkowski et al. (1992) (KPSS) have been applied. All tests are consistent with the presence of a unit root in the three price series, satisfying the first necessary condition for cointegration analyses[5].

3.2. Cointegration analysis

In this section we address the first step to specify a TVECM (i.e. testing for cointegration and estimating the cointegrating relationship). Cointegration is tested using the likelihood ratio test introduced by Johansen (1988). Escribano and Mira (1996) show that the cointegrating vector can still be estimated superconsistently in the presence of neglected non-linearity in the adjustment process. Before determining the cointegration rank, each system has to be correctly specified. More precisely, what deterministic components must be included and what is the optimum lag that ensures that residuals are approximately white noise and have zero autocorrelations at all lags. In this paper, the optimum lag has been selected on the basis of the Akaike Information Criterion (AIC) and the Likelihood Ratio test proposed by Tiao and Box (1981). Both tests provide consistent results and indicate that four lags would be the optimum lag in the system.

Misspecification tests for autocorrelation and normality, described in Doornik and Hendry (1997), have been carried out for each system to check for the statistical adequacy of the model. Results indicate that models specified above are quite satisfactory (Table 1). However, due to excess kurtosis, normality of residuals is rejected which may be caused by neglected nonlinearity. Table 1 also shows the results of the Johansen likelihood ratio tests for cointegration rank. At the 5% level of significance, both tests indicate that the null hypothesis of one cointegrating vector cannot be rejected. Given that the cointegrating rank is one, we have tested whether the price transmission between farm and retail prices is perfect in the long run. This hypothesis states that the cointegrating vector β should satisfy the long-run price homogeneity condition (1,-1). All restriction tests on the cointegrating vector

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[4] The procedure is similar to that used by Franses (1991), for monthly data, and it is based on the decomposition of the polynomial ($1-L^{52}$). The description of the procedure has not been included due to space limitations. In any case, results are available from authors upon request.
[5] Results are not shown due to space limitations. They are available upon request.

are asymptotically $\chi^2(v)$ distributed where v is the number of imposed restrictions[6]. Results from the Likelihood Ratio (LR) statistic (second row of Table 1) show that the homogeneity restriction cannot be rejected and has empirical support.. The restricted cointegrating vector is given by:

$$LnRP - lnFP = 0.448 \quad (10)$$

The constant term in (10) represents the price spread at the retail levels. Taking into account that all prices are expressed in logarithms, equation (10) represents percentage spread models with a mark-up of (e^{α}-1) (with α being the constant) (Tiffin and Dawson, 2000). Hence, the retail marketing margin can be expressed as follows:

$$\text{Retail margin} = (e^{\alpha}-1)\times FP\times 100 = 56\%FP \quad (11)$$

3.3. Threshold cointegration

Once the presence of a long run equilibrium relationship between the two prices has been detected the next question is whether possible nonlinearities exist in the adjustment process. This question will be analysed using the procedure described in Section 3. We start by testing nonlinearity and, in case the null of linearity is rejected, the number of regimes in the TVECM is determined considering the estimated cointegrating vector, given in (11), as the threshold variable (ω_{t-1})[7]. Results from the LR linearity test against the alternative of a multivariate $TVECM_3$ ($LR_{1,3}$) are shown in Table 2 and indicate that the null of linearity is rejected at the 5% level, in favour of the threshold model.

Table 1. Cointegration analysis in FP-WP and WP-RP systems[a]

Cointegration Trace Statistic		H_0: r=0	H_0: r=1
		76.08	7.83
	Critical Value (5%)	20.12	9.17
H_0:β=(1,-1)		3.05	(0.08)
Multivariate misspecification tests	A_LM(1)[b] = 7.33 (0.11)		
	A_LM(52)[c] = 8.23 (0.08)		
	N_LM = 14.85 (0.00)		

Values in parentheses are p-values

A_LM(i) is the Godfrey multivariate test for autocorrelation of order i.

N_LM is the Doornik and Hendry (1997) multivariate test for normality.

Table 2. Tests for nonlinearities in price adjustments[a,b]

	LR_{13}	LR_{23}
Test statistic	63.28	49.82
FR critical value (5%)[c]	29.17	28.91
PR critical value (5%)[d]	32.76	30.17
Threshold parameters	$\hat{\lambda} = (-0.028, 0.054)$	

a The $LR_{1,3}$ tests the null of linearity against the alternative of a TVECM (Lo and Zivot, 2001).

b The $LR_{2,3}$ tests the null of a two-regime TVECM against the alternative of three-regime TVECM (Lo and Zivot, 2001).

c Critical values are obtained using the fixed regressor (FR) bootstrapping technique (Hansen, 1999; and Hansen and Seo, 2002).

d Critical values are obtained using the parametric residual (PR) bootstrap algorithm Hansen, 1999; and Hansen and Seo, 2002).

Given that linearity is rejected in favour of threshold nonlinearity, next we test which threshold model is more appropriate to characterize the nonlinear dynamic adjustments of prices using the $LR_{2,3}$ statistic given in expression (6). As can be observed from Table 2, the LR statistic reject the null of a $TVECM_2$ against the alternative of a three-regime $TVECM_3$, suggesting that price transmission along

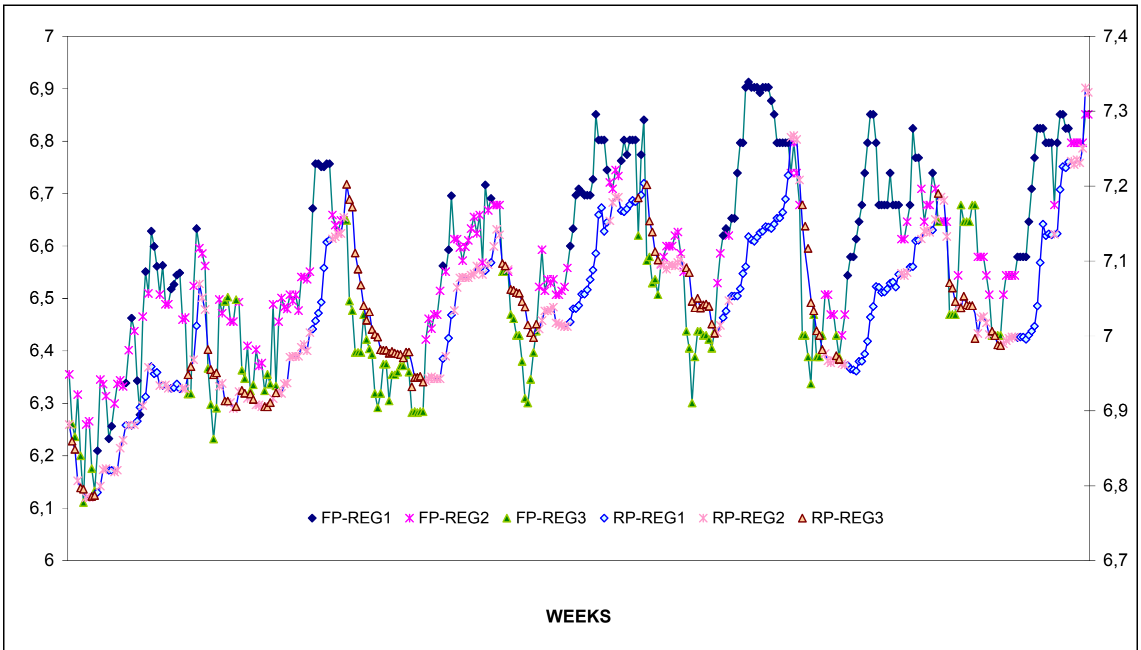
[6] For further details, see Johansen and Juselius (1994) and Johansen (1995).

[7] The residuals obtained from equations (12) and (13) can be interpreted as deviations from a long-term equilibrium.

the Spanish lamb marketing chain can be characterised by a three-regime threshold process. At the bottom of Table 2 the estimated threshold parameter from the $TVECM_3$ is showed ($\hat{\lambda} = (-0.028, 0.054)$). In other words, and taking into account (11), the TVECM splits the price adjustment processes depending on whether the retail marketing margin lies below 52%, above 71% or between 52% and 71%. Figure 1 reports the evolution of farm and retail prices under the three regimes according to the corresponding threshold parameter. As can be observed from Figure 1, the first regime (marketing margin below the threshold level) can be associated with increasing phases of lamb prices (excess demand), while the third regime (marketing margin above the threshold level) seems to be associated with periods of declining prices (excess supply). The second regime is associated with transition periods, that is, when prices star to rise or to decline. These results are quite consistent with those found for the United States by Breimyer (1957) who concluded that, in the short-run, marketing margins tended to increase when production also increases, while they decreased with production shortages, being quite stable in the long-run.

The estimated $TVECM_3$ coefficients are shown in Table 3 as well as results from misspecification tests. As can be observed, results of diagnostic tests suggest that the estimated models are adequate as there is no evidence for remaining residual autocorrelation, ARCH tests fail to reject the null of homocedasticity and, finally, normality cannot be rejected. Moreover, the estimated parameters in the outer regimes are significant and have the expected sign. However, in the middle regime (regime 2) adjustment coefficients are not significant, indicating that adjustment only takes place till the edge of the threshold band. Within the band, the two prices move closer to each other but without following any specific pattern.

Figure 1. Classification of prices evolution under the three regimes (prices in logs)



Considering this result, the $TVECM_3$ could be re-specified as a Band-TVECM as it has been defined in Section 3. A Wald test is carried out to check if adjustment coefficients in the middle regime are jointly significant. Results indicate that the null of no significance cannot be rejected at the 5% significance level (the Wald statistic is 3.31, while the critical value is 5.99 at 5% of significance level). Consequently, it can be concluded that Band-TVECM is more appropriate than the unrestricted TVECM to represent the asymmetric adjustments of lamb prices along the marketing channel.

The estimated parameters of the Band-TVECM are given in Table 4. Furthermore, we include the estimates of the adjustment parameters $\hat{\rho}_i$, which measure how the cointegrating vector reverts to the regime-specific mean (see footnote 2). As can be observed, the estimated parameters $\hat{\rho}_i$ in regime 1 are always lower than those in the upper regime. A smaller $\hat{\rho}_i$ means that price adjustments after disequilibria are faster. In the lower regime $\hat{\rho}$ is 0.682 and increase to 0.867 in the upper regime.

Table 3. Estimated parameters of the $TVECM_3$[a]

	Regime 1[b] $\omega_{t-1}(\beta_2) < -0.028$	Regime 2[b] $-0.028 \le \omega_{t-1}(\beta_2) \le 0.054$	Regime 3[b] $\omega_{t-1}(\beta_2) > 0.054$
$\begin{pmatrix} \alpha_1^i \\ \alpha_2^i \end{pmatrix}$	$\begin{pmatrix} \underset{(0.056)}{-0.18} \\ \underset{(0.022)}{0.052} \end{pmatrix}$	$\begin{pmatrix} \underset{(0.07)}{0.002} \\ \underset{(0.015)}{-0.013} \end{pmatrix}$	$\begin{pmatrix} \underset{(0.020)}{-0.048} \\ \underset{(0.063)}{0.102} \end{pmatrix}$
Misspecification tests			
BG(1)-FP[c]	2.59	BG(1)-RP	0.44
BG(52)-FP[c]	1.46	BG(52)-RP	1.13
ARCH(1)-FP[c]	3.84	ARCH(1)-RP	3.32
ARCH(52)-FP[d]	3.76	ARCH(52)-RP	3.86
JB-FP[e]	3.04	JB-RP	4.02
% of observations	33.33	38.33	28.33

a. Values in parentheses are standard deviations.

b. $\omega_{t-1}(\hat{\beta}_2) = RP - WP - 0.448$.

c. BG(i) is the Breush-Godfrey test for autocorrelation of order i (Critical value at the 5% level of significance is 3.84).

d. ARCH (i) is the Engle test for conditional heteroscedasticity of order I (Critical value at the 5% level is 3.84).

e. JB is the Jarque-Bera test for normality. Critical value at the 5% level of significance is 5.99

The speed of adjustment is usually measured by the so-called half-life [ln(0.5)/ln($\hat{\rho}_i$)] which states the number of periods required to reduce one-half of a deviation from the long-run equilibrium (Obstfeld and Taylor, 1997). Taking into account the results mentioned in the above paragraph, the half-life increases from 1.80 weeks to 4.83 weeks. This results indicates that the adjustment induced by a negative deviation from the stationary price relationship is much faster than when it is induced by a positive deviation.

In any case, as we have already mentioned in the previous section, the key feature in threshold models is the pattern of the estimated coefficients of the α matrix (α_{ij}) associated to the cointegrating vector $\omega_{t-1}(\beta)$. These coefficients can be useful to analyse which prices "equilibrium adjust", and which do not. The first interesting point to note is that the estimated coefficients corresponding to the lower regime, in absolute values, are larger than those corresponding to the upper regime, indicating that the speed of adjustment is more rapid for negative than for positive deviations from the threshold values. Given that the lower (upper) regime indicates that the marketing margin is below (above) its long-run equilibrium value, this suggests that prices react more rapidly when the margin is squeezed than when it is stretched. These results would appear to be quite consistent with those reported by von Cramon-Taubadel (1998).

During the lower-margin regime (first regime), the adjustment coefficients are significant, indicating a feedback effect between the two prices. In addition, estimated coefficients indicate that the speed of adjustment of the retail prices is more rapid than that of the farm prices (after a negative deviation from the long-run equilibrium relationship, the retail price adjusts by eliminating 21% of such a negative impact generated in the previous period, while in the case of the farm price the adjustment is only about 10.5%). In the upper regime, adjustment coefficients are significant for the farm price, but not for the retail price. Thus, a positive shock on the price spread between the two levels of the marketing chain will initiate an adjustment process in the farm price, but not in the retail price, indicating that retail prices are sticky relative to farm prices when the marketing margin is squeezed.

Table 4. Estimated parameters of the Band-TVECM[a]

Regime 1[b]			Regime 3[b]		
$\begin{pmatrix}\alpha_1^i \\ \alpha_2^i\end{pmatrix}$	ρ^c	Half-Life[d]	$\begin{pmatrix}\alpha_1^i \\ \alpha_2^i\end{pmatrix}$	ρ^c	Half-Life[d]
$\begin{pmatrix}\underset{(0.032)}{-0.213} \\ \underset{(0.014)}{0.105}\end{pmatrix}$	0.682	1.80	$\begin{pmatrix}\underset{(0.043)}{-0.052} \\ \underset{(0.023)}{0.087}\end{pmatrix}$	0.867	4.83

a. Values in parentheses are standard deviations
b. Regimes 1 and 3 have been already defined in Table 3.
c. ρ is the adjustment coefficient which measures how the cointegrating vector reverts to the regime-specific mean (see footnote 2 for its mathematical expression).
d. Half Life is defined as [ln(0.5)/ln($\hat{\rho}_i$)]

3.4. Short-Run Dynamics

Short-run dynamics have been analysed by computing the IRF, which show the response of each price in the system to a shock in any other price. In this study, Non linear IRF (NIRF) have been calculated for each regime prices are expected to react (regimes 1 and 3). In a context of non-linear models, NIRF are a very useful tool, as they allow us to differentiate responses to both positive and negative shocks. Moreover, the time at which the shock takes place is relevant, and thus, we could expect different responses depending on which of the regimes the shock is produced.

In order to analyse the asymmetric behaviour of price adjustments, the NIRF have been computed for $\delta=\pm 1$ and $\pm$ 2 and for history-specific regimes such that the long-run equilibrium relationship [$\omega_t(\beta_i) = \beta' P_{t-1}$] (i=1,2 for the first and second system, respectively) is above or below the upper and lower threshold values. In each regime, the NIRF for each forecasting horizon is the average across all possible N_i histories (with N_i being the number of observations in the ith regime). For each response, we have computed the corresponding 95% confidence intervals using bootstrapping techniques based on 5,000 replications[8]. Figure 2 shows main results. Under the first regime, i.e. when prices are increasing (Figure 2, Panel a), a 1% positive shock to the retail price generates and immediate and significant response of both prices. However, the magnitude of such responses is quite different. The farm price exhibits a certain delay in adjusting to the new situation, reaching the maximum response after three weeks. Thus, although in the long-run both prices are perfectly integrated, in the very short-run retailers benefit from a demand shock as the price spread increases by 50%. The situation is quite similar when the magnitude of the shock is 2%, generating responses, which, approximately, doubled those generated by a 1% positive shock.

Responses of retail and farms prices to a negative demand shock at the retail level have a similar pattern than in the case of positive shocks although two main differences exist. First, responses are significant for a shorter period and, second, the magnitude of the response is lower during the first 10 weeks after the shock, mainly in the case of the retail price, suggesting that positive shocks are more persistent and generate positive asymmetries. Moreover, although in the first week the negative response of the retail price in higher than that of the farm price, the situation reverse from then generating increasing price spreads since the third week after the shock.

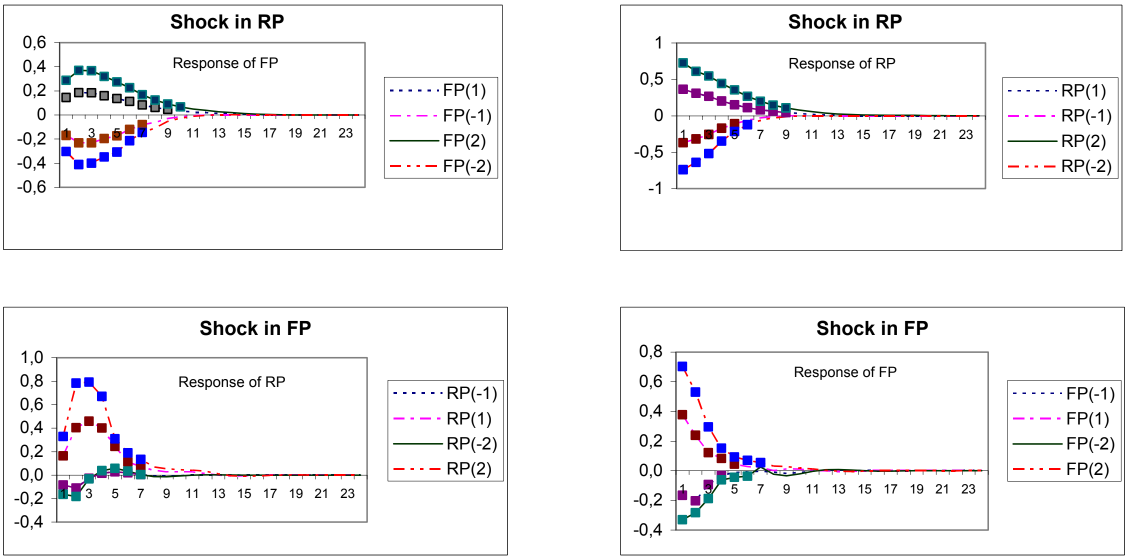
A positive shock in the farm market notably stretches the marketing margin the first week after the shock, as the farm price response is about 40% of the magnitude of the initial shock while in the case of the retail price, the response is only about 20%. However, during the following weeks, the retail price overreacts to the initial shocks, increasing the price spreads for about 6 weeks after the

[8] All analyses have been carried out in GAUSS. We are grateful to Dr. van Dijk for providing valuable information on haw to tackle this cumbersome task.

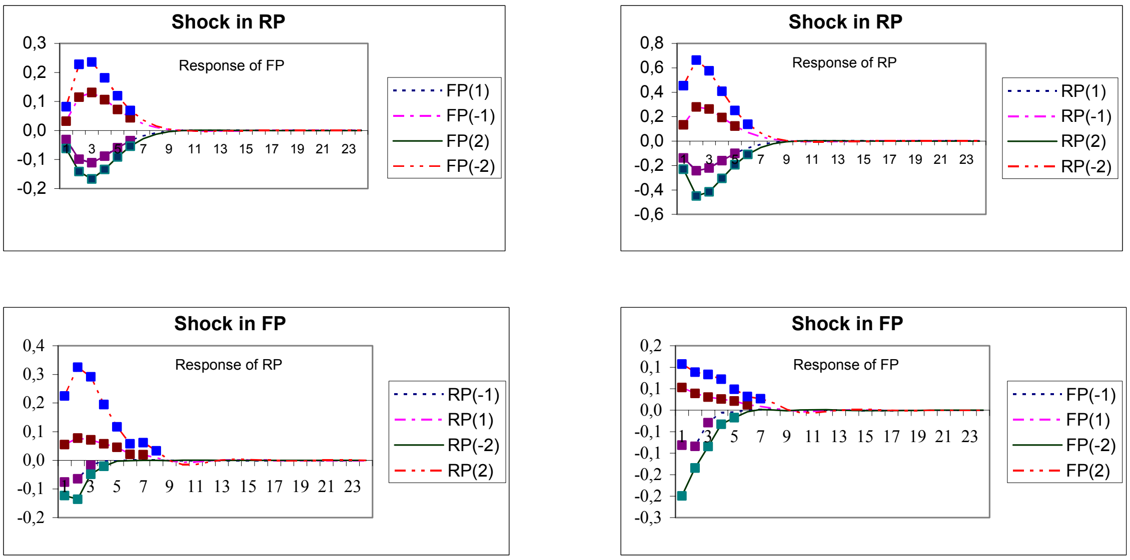
shock. The existence of only one week of delay to react has to do with the specific characteristics of lamb. It is a perishable product mainly sold in big pieces the butcher has to cut. No labels, apart from specific quality labels, are present. Thus, menu costs are irrelevant, as retailers have to change only the price. In the case of specific cuts already packed, the stock disappears in less than one week. If the magnitude of the initial shock doubles, then the magnitude of the responses is more than proportional.

Figure 2. Impulse response functions to a 1% and 2% positive and negative shock for system WP-RP under the two regimes

Panel a) Regime 1 ($\omega_{t-1}(\beta_2) < -0.028$)



Panel b) Regime 3 ($\omega_{t-1}(\beta_2) > 0.054$)



Note that squares indicate that the response is significant at the 5% level

Responses of farm and retail prices to excess supply shocks (independently of the magnitude) have a similar path. However, the magnitude of such responses are different, being persistently higher in the case of the farm price, thus generating increasing price spreads, which benefit retailers. As can easily be observed, comparing the responses to positive and negative farm market shocks, the price adjustment process is positive-asymmetric (price increases are transmitted faster than price decreases). Finally, the magnitude of the asymmetric effect is greater in the case of the retail price, suggesting that inflation in food products is not generated by cost increases, but rather by increases in marketing margins. These results seem to indicate that retailers have certain market power in the lamb market in Spain, as is the case with most perishable products. As Bettendorf and Verboven (2000) show, price behaviour is related to market concentration and oligopsonistic behaviour. In fact, retailers are much more concentrated than farmers, at least in the case of supermarkets and hypermarkets chains operating at national level.

Under the third regime, i.e. when prices are falling (Figure 2, Panel b) the general pattern are more or less the same although three main differences may be appreciated. First, the magnitude of the responses is lower, especially in the case of negative shocks. Second, the convergence towards the long-run equilibrium takes place more quickly, independently of the magnitude of the initial shock. Third, when declining prices variations in the magnitude of the initial shock generate responses more than proportional when the shock is positive but much less than proportional when the shock is negative.

In general terms, all considered cases lead to increasing price spreads in the short run, benefiting retailers, with the only exception of a negative shock in the retail price, in which retail prices decrease slightly faster than farm prices. In an environment of declining prices, retailers are not able to push farm prices significantly down in order to guarantee long run supply. In any case, the reduction of the price spread in this specific case is substantially lower in absolute values than the increase that takes place after a positive shock. Finally, short-run responses to positive shocks are higher than those for negative shocks, indicating, as in the first regime, the existence of positive asymmetries.

4. CONCLUSIONS

This paper has explored the non-linearity in the price transmission mechanism along the lamb marketing chain in Spain. The methodology used has been based on the specification and estimation of a three-regime TVECM in which regimes are associated with price cycles. Moreover, price reactions in the intermediate regime are not significant allowing us to specify a Band-TVECM. Obtained results suggest a number of points.

In the long-run, prices at both extremes of the marketing chain are perfectly integrated; that is to say, any change in any of the prices is fully transmitted to the other. However, in the short-run, price behaviour has to be with the structure of the retail sector. Retailers have clearly market power. Two thirds of total lamb sales at retail level are located in supermarkets and hypermarkets in which the market share of the top-five is around 60%. The main conclusion is that, in an environment of increasing prices, retailers benefit from any shock, whether positive or negative, that affects supply or demand conditions. In the first case, a shock to the farm price notably stretches the marketing margin in the very short run (one week after the shock) but then the retail price overreacts to the initial shock, increasing price spreads for about one month. In the second case, the price spread immediately increases by 50% and persists during one month and a half. Price adjustments are positive-asymmetric suggesting that retail prices show more nominal flexibility when they are increasing.

Under a price-declining situation, the general pattern is rather similar. However, responses converge more rapidly to the long-run equilibrium and they are much higher when shocks are positive than when they are negative as further price decreases can generate short-run losses.

The analysis has focused on vertical price adjustments in the Spanish lamb marketing chain. It can be extended in several directions. First, a natural extension will be to investigate other meat sectors in Spain with different market structures (different degrees of market integration) or other food sectors with different characteristics (branded products, more processed products, non-perishable products, etc) to better understand the price transmission mechanism and to what extent farm prices are responsible of inflation. Also, further applications to the same sector in other countries with different market structures would allow us to link our results with market power or holding stocks policy. Finally, further refinements from the methodological point of view could be used in the future as new theoretical econometric issues arise in the context of non-linear models in a multivariate framework.

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