Price transmission asymmetries in the Spanish lamb sector

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Abstract

The analysis of asymmetries in the price transmission mechanism at different levels of the marketing chain provides some interesting information about the degree of competition in vertical related markets. The objective of this paper is to investigate the non-linear adjustments of prices along the lamb sector in Spain. The methodology used is based on the multivariate approach to specify and estimate a Threshold Autoregressive Model. Price relationships at farm, wholesale and retail levels are considered. Results indicate that in the long-run price transmission is perfect and any supply or demand shocks are fully transmitted to all prices in the system. In the short-run, analyses suggest that the high degree of horizontal concentration among retailers allow them to have market power. Responses to any shock generate an increase of the retail price spread which is more evident when prices show an upward trend.

Key words: Price transmission, asymmetries, TAR models, lamb, Spain

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1 INTRODUCTION

The issue of asymmetries in price relationships along the food chain has been one of the main research interests among agricultural economists, given that such asymmetries are a good indicator of market performance. In this way, it is possible to determine whether the lagged adjustment processes at different levels of the marketing chain, after a price change, for instance, at the farm level, are the same for price increases as for price decreases. It is assumed that symmetric relationships are representative of competitive markets, while asymmetric responses are linked with the existence of some market imperfections[1]. Using variations of a model first developed by Wolframm (1971) and later modified by Houck (1977), most authors have found evidence of both asymmetries in price adjustments and a cost-push price transmission mechanism for different products[2] (see, for instance, Ward, 1982; Kinnucan and Forker, 1987; Hahn, 1990; and Hansmire and Willett, 1992; among others).

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However, the empirical models used by the above authors to investigate asymmetries in price transmission have been criticised for the following reasons: i) these models have been used without adequately analysing the time series property of the 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), and ii) 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 misspecificatioon 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).

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The second limitation is that, generally, it is assumed that the underlying price transmission mechanism is linear. However, 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. However, these reactions may also be different for positive or negative shocks. If this is the case, then threshold models of dynamic economic equilibrium are more appropriate when analysing dynamic price relationships between markets along the food chain. From amongst these, the Threshold Autoregressive Models (TAR) have become increasingly popular.

In this context, two different methodological approaches have been developed to analyse cointegration relationships with asymmetric deviations. The first 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 (1999). In a similar way to the two-step Engle and

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[1] Ward (1982), Kinnucan and Forker (1987) and Bailey and Brorsen (1989) provide a number of theoretical reasons for asymmetries in price adjustments.

[2] The only exception is Boyd and Brorsen (1988), who do not find asymmetric price relationships in the US pork sector.

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. However, it does not consider the threshold behaviour in the border bivariate model for log prices. Furthermore, the univariate process assumes that one of the two prices is exogenous (for instance, that changes in consumer prices are caused by changes in prices at a lower level, say farm prices, in the marketing chain). Goodwin and Holt (1999) applied such an approach when studying the price transmission and asymmetric adjustment in the US beef sector.

The second approach has been suggested by Hansen and Seo (2001). As in Balky-Fomby (1997), the Hansen and Seo model is a vector error correction model (VECM) with one cointegrating vector and a threshold effect based on the error-correction term. However, unlike the first approach, which is based on univariate estimation and testing methods, the estimation and tests statistics proposed by Hansen and Seo (2001) are for the complete multivariate threshold model. 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 we have adopted in this paper.

Against this background, we apply the Hansen-Seo (2001) approach to analyse the price transmission mechanism among farm, wholesale and retail marketing channels in the Spanish lamb markets. Particularly, we try to answer to the following three questions: i) whether there is a perfect transmission mechanism in the long run between prices along the lamb marketing channels, ii) whether farmers benefit or not from unanticipated supply or demand shocks, and finally iii) if price adjustment processes are symmetric or asymmetric in the Spanish lamb sector. To answer the last question, we consider reactions in lamb prices at different level of the marketing chain to both positive and negative supply and demand shocks. The extent and speed of adjustment with which shocks are transmitted among farm, wholesale and retail market prices may reflect the performance of mark participants at alternative market levels.

The rest of the paper is organised as follows. In section two we provide a brief 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

2.1 *Threshold cointegration*

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 Hansen and Seo (2001), a two-regime threshold vector error correction model (TVECM), can be written as:

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

where $\omega_t(\beta) = \beta' P_{t-1}$ is the threshold variable which is equal to the residual from the cointegrating relationship, $\omega_t(\beta) = \beta' P_{t-1} = P_{1t} - \beta_2 P_{2t}$, and λ is the threshold parameter.

The model can be alternatively written as:

$$\Delta P_t = I_t(\beta,\lambda) A_1' X_{t-1}(\beta) + [1 - I_t(\beta,\lambda)] A_2' X_{t-1}(\beta) + \varepsilon_t \tag{3}$$

where:

$I_{t(\beta,\lambda)}$ is a heavyside indicator function such that: $I_t(\beta,\lambda) = \begin{cases} 1 \text{ si } \omega_{t\text{-}1}(\beta) \le \lambda \\ 0 \text{ si } \omega_{t\text{-}1}(\beta) > \lambda \end{cases}$

$$X_{t-1}(\beta) = \begin{pmatrix} \omega_{t-1}(\beta) \\ \Delta P_{t-1} \\ \Delta P_{t-2} \\ \vdots \\ \Delta P_{t-k+1} \end{pmatrix} \text{ and } A^i = \begin{pmatrix} \alpha^i \\ \Gamma_1^i \\ \vdots \\ \Gamma_{k-1}^i \end{pmatrix}$$

A_1 and A_2 are the parameters vectors associated to the regime 1 [$\omega_{t\text{-}1}(\beta) \le \lambda$] and 2 [$\omega_{t\text{-}1}(\beta) > \lambda$], respectively.

As can be observed, in the TVECM specified in (3) the dynamic behaviour and the adjustment towards the long-run equilibrium relationship [$\omega_t(\beta) = \beta' P_{t-1}$] depend on whether prices are above or below their long-run equilibrium value.

Note that when the long-run parameter β and the threshold value (λ) are both fixed (known a priori)[3], the model is linear in the remaining parameters. In such circumstance and under the assumption that errors ε_t are iid gaussian, the parameters in model (2) can be estimated by OLS, where the estimated parameters are given by:

$$\hat{A}_1(\beta,\lambda) = \left(\sum_{t=1}^{n} X_{t-1}(\beta) X_{t-1}'(\beta) I_t(\beta,\lambda) \right)^{-1} \left(\sum_{t=1}^{n} X_{t-1}(\beta) \Delta P_t I_t(\beta,\lambda) \right) \tag{4}$$

[3] In a number of economic applications it seems consistent to set a priori the threshold value as equal to zero, so that the cointegrating vector coincides with the attractor (Abdulai, 2000).

$$\hat{A}_2(\beta,\lambda) = \left(\sum_{t=1}^{n} X_{t-1}(\beta)X'_{t-1}(\beta)[1-I_t(\beta,\lambda)]\right)^{-1}\left(\sum_{t=1}^{n} X_{t-1}(\beta)\Delta P_t[1-I_t(\beta,\lambda)]\right) \quad (5)$$

$$\hat{\Sigma}(\beta,\lambda) = \frac{1}{n}\sum_{t=1}^{n}\hat{\varepsilon}_t(\beta,\lambda)\hat{\varepsilon}_t(\beta,\lambda)' \quad (6)$$

where $\hat{\Sigma}(\beta,\lambda)$ is the estimated covariance matrix of the model (3).

However, in general, the value of λ is unknown and needs to be estimated along with the remaining parameters of the model. Hansen and Seo (2001) provide a search procedure to jointly estimate the values of the two-dimensional space (β,λ) which consists of the following steps. First, form a grid on $\lambda \in [\lambda_L, \lambda_U]$ and $\beta \in [\beta_L, \beta_U]$. In the first case, the particular choice for the interval $[\lambda_L, \lambda_U]$ is based on the restriction that each regime contains at least a pre-specified fraction (π_0) of the total sample (T), that is to say[4]:

$$\pi_0 \le T^{-1}\sum_{t=1}^{T}[I_t(\beta,\lambda) \le \lambda] \le 1-\pi_0 \quad (7)$$

In the case of the cointegrating vector β, the interval of allowable cointegrating parameters can be formulated taking into account the estimate of β from the linear VECM in (1) (dented as $\tilde{\beta}$).

For each value of (λ,β) on this grid, $\hat{A}_1(\beta,\lambda)$, $\hat{A}_2(\beta,\lambda)$, and $\hat{\Sigma}(\beta,\lambda)$ are calculated as defined in (4), (5) and (6), respectively. The values of λ and β on this grid which yield the lowest value of $\log\left|\hat{\Sigma}(\beta,\lambda)\right|$ are taken as the final estimates of both types of parameters, then:

$$(\hat{\beta},\hat{\lambda}) = \underset{\lambda\in[\lambda_L,\lambda_U]\,y\,\beta[\beta_L,\beta_U]}{\arg\min}\left(\log\left|\hat{\Sigma}(\beta,\lambda)\right|\right) \quad (8)$$

Once the parameters of model (3) have been estimated, the next step is to test if the dynamic behaviour and the adjustment towards the long-run equilibrium relationship is linear or exhibits threshold non-linearity. This hypothesis can be formulated as:

H_0: $A_1 = A_2$ (symmetric adjustment)

against the alternative

H_a: $A_1 \neq A_2$ (asymmetric adjustment)

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 liner model. As a consequence, the conventional statistical theory cannot be applied

[4] The particular choice of π_0 is somewhat arbitrary. In any event, each regime needs to have sufficient observations to adequately identify the regression parameters.

to obtain the asymptotic distribution of the statistics (see Davies, 1987; Hansen, 1996 and Hansen and Seo, 2001). Given that the test statistic has a non-standard distribution, the critical values have to be determined by simulation methods such as the bootstrapping technique (for more details, see Hansen, 1997 and Hansen and Seo, 2001). As a solution to the above-mentioned problem, Hansen and Seo (2001) propose the following Sup-LM statistic based on the Lagrange Multiplier (LM) Principle:

$$\mathrm{SupLM} = \sup_{\lambda_L \leq \lambda \leq \lambda_U} \mathrm{LM}(\tilde{\beta}, \lambda) \tag{9}$$

where $\mathrm{LM}(\tilde{\beta},\lambda)$ is the heteroskedasticity-robust Lagrange Multiplier (LM) statistic, which tests the restriction as given by the null hypothesis.

The SupLM is equivalent to the supremum over the set $\lambda \in [\lambda_L, \lambda_U]$. As the distribution of the Sup-LM statistic is non-standard, Hansen and Seo (2001) suggest using the fixed regressor bootstrap or, alternatively, a parametric residual bootstrap algorithm, to compute the p-value for the linearity tests.

2.2 Non-linear impulse response functions

Once the TVEC 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 no- linear models[5]: 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:

$$\begin{aligned}\mathrm{NIRF}(n,\delta,\varphi_{t-1}) = {} & E[P_{t+n} \mid \varepsilon_t = \delta, \varepsilon_{t+1} = \ldots = \varepsilon_{t+n} = 0, \varphi_{t-1}] \\ & - E[P_{t+n} \mid \varepsilon_t = 0, \varepsilon_{t+1} = \ldots = \varepsilon_{t+n} = 0, \varphi_{t-1}] \text{ for } n = 0,1,\ldots N\end{aligned} \tag{10}$$

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:

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[5] In the linear model, IRF are symmetric, in the sense that a shock of size $-\delta$ has exactly the opposite effect to that of a shock of size δ.

$$\mathrm{NIRF}(n,\varepsilon_t,\Omega_{t-1}) = E[P_{t+n} \mid \varepsilon_t,\Omega_{t-1}] - E[P_{t+n} \mid \Omega_{t-1}] \quad (11)$$

From (11), 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 (when the series are increasing or decreasing), and under different sizes and signs of the initial shock. In particular, the NIRF can be used to evaluate the degree of asymmetric responses over time. In this sense, Potter (1995) defines a measure to assess the asymmetric response to a particular shock, given a particular history φ_{t-1}, as the sum of NIRF for this particular shock and the NIRF for the shock of the same magnitude but with opposite sign, that is to say[6]:

$$\mathrm{ASY}(n,\varepsilon_t,\varphi_{t-1}) = \mathrm{NIRF}(n,+\delta_i,\varphi_{t-1}) + \mathrm{NIRF}(n,-\delta_i,\varphi_{t-1}) \quad (12)$$

3. EMPIRICAL ANALYSIS

As mentioned in the introduction, in this paper we apply the Hansen-Seo approach to analyse the price transmission mechanism along the Spanish lamb marketing chain. The methodological approach consists of the following steps. After testing for unit roots, the number of existing cointegrating relationships among all non-stationary variables is tested using the Johansen (1988) procedure. Second, taking into account the results from the previous step, several restrictions are imposed on the cointegrating vectors in order to test for long-run prices homogeneity. Once the long-run behaviour is analysed, the Hansen-Seo approach is used to determine whether or not threshold behaviour in the error correction term can be rejected. Finally, if the price transmission mechanism follows a two-regime bivariate threshold error correction model (TVECM), then non-linear Generalised Impulse response functions are calculated in order to analysis the response of each prices to unanticipated positive and negative shocks.

In this paper we consider farm (FP), wholesale (WP) and retail (RP) lamb prices. Weekly data from 1993:1 to 1999:52 are used. Farm and wholesale prices are taken directly from Spanish Ministry of Agriculture, Food and Fisheries (MAPA)[7]. Retail prices are taken from the Boletín Económico del ICE (Ministry of Finance). 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, pp. 31-32). 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.

Before implementing the Johansen and Juselius procedure for the cointegration analysis among the price series, we first examine the stochastic time series properties. The order of

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[6] For more details, see Franses and van Dijk (2000)

[7] The farm price is a weighted average of the five more representative markets in Spain (Ebro, Talavera de la Reina, Zafra, Albacete and Medina del Campo). The wholesale price is given by the price in Mercamadrid, the most important wholesale market for meat in Spain. Both prices are expressed in pesetas per kilogram of carcass weight. Data are not published, but these prices are those that are sent to Brussels in order to obtain the European representative market price for lamb to calculate the ewe premium. We are grateful to the Spain Ministry of Agriculture for allowing us to use such data.

integration of each series is calculated based on the Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979, 1981) and Phillips-Perron (PP) (Phillips and Perron, 1988) tests. The unit root tests are conducted for each price series following a sequential procedure explained in Harris (1995), in which special attention is paid to the lags and deterministic components included in the model[8]. Both tests indicate that the three price series are clearly I(1) processes.

3.1. Cointegration analysis

Given that all price series under study are integrated of the same order, the Johansen procedure is used to test for cointegration among the time series. As a first step in the analysis, the unrestricted VAR is specified including three lags (based on results from the Tiao and Box (1981) Likelihood ratio test) and a restricted constant term lying in the cointegration space in (1), indicating that there are no linear deterministic trends in the data[9]. Before applying the reduced rank tests, the multivariate test for autocorrelation (Godfrey, 1988) has been carried out to check for the statistical adequacy of the model. The results indicate that the model defined above can be considered as being correctly specified (6.95 and 9.44 for autocorrelation of order 1 and 52, respectively, with a critical value of 16.9 at the 5 per cent level of significance).

Table 1 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 two cointegrating vectors cannot be rejected, suggesting that there is a common trend driving the three prices. The existence of two cointegrating vectors indicates that any variable may be cointegrated in terms of any of the other variables. However, cointegration as such does not say anything about the direction of causality. Moreover, the existence of cointegration by itself does not imply which prices "equilibrium adjust" and which do not, and which price can be a price leader. Information about such a feature can be provided by the α_{ij} coefficients. We will return to this point later in this paper. A more relevant issue related to this study consists of testing whether the price transmission between farm and wholesale and between wholesale and retail prices is perfect in the long run. This hypothesis states that each cointegrating vector should satisfy the long-run price homogeneity condition (1,-1), such that the matrix β adopts the following expression:

$$\beta' = \left(FP, \quad WP, \quad RP, \quad c\right) \times \begin{pmatrix} -1 & 0 \\ 1 & -1 \\ 0 & 1 \\ * & * \end{pmatrix} \tag{13}$$

The Johansen and Juselius (1994) procedure has been used to test for restrictions in (10)[10]. The Likelihood Ratio (LR) statistic of two over-restrictions was 5.67, which is well under the critical value at the 5% the level of significance ($\chi^2(2)$=5,99). Thus, the homogeneity restrictions cannot be rejected and have empirical support. Consequently, it can be concluded that in the long run any change in any of the prices at different levels of the Spanish lamb marketing chain is fully transmitted to the rest. The restricted cointegrating vectors are given by:

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[8] Results are not shown due to space limitations. They are available upon request.

[9] This specification was consistent with the results from unit root tests, which indicated that $E[\Delta P_t]=0$ for all price series, implying that some equilibrium means were different from zero.

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

$$\text{LnRC} - \ln\text{WM} = 0.448 \quad (14)$$

$$\text{LnWP} - \ln\text{FP} = 0.157 \quad (15)$$

Table 1. Results from multivariate cointegration rank tests

λ-max	Trace	H_0: r	P - r	λ-max CV(95%)	Trace CV(95%)
55.23	84.07	0	3	22.00	34.91
20.42	28.84	1	2	15.67	19.96
8.43	8.43	2	1	9.24	9.24

The constant terms in (14) and (15) represent the price spread at the retail and wholesale levels, respectively. Taking into account that all prices are expressed in logarithms, (11) and (12) represent percentage spread models with a mark-up of (e^{α}-1) (with α being the constant) (Tiffin and Dawson, 2000). Hence, the wholesale and retail marketing margins can be expressed as follows:

$$\text{Wholesale margin} = (e^{\alpha}-1)\times\text{FP}\times 100 = 16\%\text{FP} \quad (16)$$

$$\text{Retail margin} = (e^{\alpha}-1)\times\text{WP}\times 100 = 52\%\text{WP} \quad (17)$$

3.2. Threshold cointegration

Once the cointegration relationships have been estimated, the next step consists of testing for asymmetric adjustments between the prices using the procedure described in Section 2. Nevertheless, given that the TECM defined in (3) is bivariate with only one cointegrating vector; the analysis has been carried out considering two separate subsystems. The first of those considers the existing relationship between the farm price (FP) and the wholesale price (WP), while the second analyses the existing relationship between the wholesale price (WP) and the retail price (RP).

We start the analysis by determining the lag orders in the TVECM model in each subsystem using the AIC criterion[11]. In addition, we consider the cointegrating vectors estimated in (14) and (15) as threshold variables (ω_{t-1}) in subsystems 1 and 2, respectively. The results are shown in Tables 2 and 3 for subsystems (FP-WP) and (WP-RP), respectively. As can be appreciated, the minimisation of the AIC suggests that both systems can be estimated with three lags (k-1=2). Autocorrelation tests also suggest that both models are correctly specified. In addition, Table 3 we include the results from the SupLM linearity test. In both system and for any lag order, these results show that the null hypothesis of linearity is rejected at the 5% level, in favour of the threshold model. Such results suggest that the price transmission along the Spanish lamb marketing chain can be characterised by a two-regime threshold process. The estimated TVECM for system one and two, respectively, have the following form:

[11] Following Tong (1990), the AIC for a two-regime TAR model can be defined as:

$$\text{AIC}(k_i) = T_1/2\ln(\det\Sigma_1) + T_2/2\ln(\det\Sigma_2) + 16(k_i)$$

where: T_j, j=1, 2, is the number of observations in the jth regime, and Σ_j, i=1, 2, is the covariance matrix in the jth regime

$$\begin{pmatrix} \Delta FP_t \\ \Delta WP_t \end{pmatrix} = \Delta P_t = \begin{cases} \sum_{i=1}^{2} \Gamma_i^1 \Delta P_{t-i} + \begin{pmatrix} \underset{(0.07)}{0.11} \\ \underset{(0.11)}{-0.25} \end{pmatrix} \omega_{t-1}(\hat{\beta}_1) + \varepsilon_t \ \text{ if } \omega_{t\text{-}1}(\hat{\beta}_1) \leq 0.057 \\ \sum_{i=1}^{2} \Gamma_i^2 \Delta P_{t-i} + \begin{pmatrix} \underset{(0.04)}{0.06} \\ \underset{(0.02)}{-0.06} \end{pmatrix} \omega_{t-1}(\hat{\beta}_1) + \varepsilon_t \ \text{ si } \omega_{t\text{-}1}(\hat{\beta}_1) > 0.057 \end{cases} \tag{18}$$

$$\begin{pmatrix} \Delta WP_t \\ \Delta RP_t \end{pmatrix} = \Delta P_t = \begin{cases} \sum_{i=1}^{2} \Gamma_i^1 \Delta P_{t-i} + \begin{pmatrix} \underset{(0.08)}{0.11} \\ \underset{(0.02)}{-0.16} \end{pmatrix} \omega_{t-1}(\hat{\beta}_2) + \varepsilon_t \ \text{ if } \omega_{t\text{-}1}(\hat{\beta}_2) \leq 0.044 \\ \sum_{i=1}^{2} \Gamma_i^2 \Delta P_{t-i} + \begin{pmatrix} \underset{(0.04)}{0.07} \\ \underset{(0.01)}{-0.03} \end{pmatrix} \omega_{t-1}(\hat{\beta}_2) + \varepsilon_t \ \text{ si } \omega_{t\text{-}1}(\hat{\beta}_2) > 0.044 \end{cases} \tag{19}$$

where: $\begin{cases} \omega_{t-1}(\hat{\beta}_1) = WP - FP - 0.157 \\ \omega_{t-1}(\hat{\beta}_2) = RP - WP - 0.448 \end{cases}$

The estimated threshold value is $\hat{\lambda}_1$= 0.057 for the subsystem (FP-WP) and $\hat{\lambda}_2$=0.044 for the subsystem (WP-RP). In other words, and taking into account (16) and (17), the TVECM splits the price adjustment processes in two regimes depending on whether the wholesale (retail) marketing margin lies above or below 23% (63%), that is to say, when, in both cases, the equilibrium marketing margin is surpassed by 7%. Although not shown due to space limitations, in the case of the subsystem (WP-RP), the first regime (marketing margin below the threshold level) can be associated with increasing phases of lamb prices (excess demand), while the second regime (marketing margin above the threshold level) seems to be associated with periods of declining prices (excess supply). However, in the case of the relationship between FP and WP (Figure 1) it is not possible to relate the two regimes to situations of excess supply or excess demand. It can only be said that the second regime takes place both at the very beginning (year 1993) and at the end of the sample (from mid 1996 on).

Returning to the estimated models given in equations (18) and (19), we can note that in both systems most of the coefficients are significant and have the expected sign. However, the key feature in equations (18) and (19) is the pattern of the estimated coefficients of the α matrix (α_{ij}) associated to the cointegrating vector $\omega_{t-1}(\beta)$, and in each regime. As we have already mentioned in the previous section, these coefficients can be useful to analyse which prices “equilibrium adjust”, and which do not. The first interesting point to note is that in both subsystems the estimated coefficients corresponding to the first regime, in absolute values, are larger that those corresponding to the second regime, indicating that the speed of adjustment is more rapid for negative than for positive deviations from the threshold values. Given that the first (second) 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 (1998).

In the FP-WP subsystem (equation 18), 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 wholesale price is more rapid than that of the farm price (after a negative deviation from the long-run equilibrium relationship, the wholesale price adjusts by eliminating 25% of such a negative impact generated in the previous period, while in the case of the farm price the adjustment is only about 11%). In the second regime, adjustment coefficients are significant for the farm price, but not for the wholesale 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 wholesale price, indicating that the wholesale price is driving the farm price when the marketing margin is squeezed.

Table 2. Lag determination, test for asymmetry and misspecificatioon tests from the TVECM for the subsystem (WP-FP)

k-1	1	2	3	4
AIC[a]	-21.60	-22.29	-22.08	-21.21
BIC[b]	-21.75	-22.06	-21.93	-21.19
SupLM test[c]	28.97	29.18	32.97	38.97
Bootstrap critical value (5%)	21.54	27.15	32.28	36.95
BG(1)-WP[d]	8.06	1.21	1.66	1.58
BG(52)-WP[d]	10.86	1.67	2.76	3.24
BG(1)-FP[d]	4.12	2.76	2.27	1.21
BG(52)-FP[d]	6.29	3.01	3.79	2.01

AIC is the Akaike Information Criteria
BIC is the Bayesian Information Criteria
SupLM is the Statistic to test for asymmetry
BG(i) is the Breush-Godfrey test for autocorrelation of order i (Critical value at the 5% level of significance is 3.84)

Table 3. Lag determination, test for asymmetry and misspecificatioon tests from the TVECM for the subsystem (RP-WP)

k-1	1	2	3	4
AIC[a]	-25.17	-25.31	-25.01	-24.85
BIC[b]	-25.03	-25.13	-24.83	-24.62
SupLM test[c]	26.12	29.232	35.97	37.14
Bootstrap critical value (5%)	21.54	27.15	32.28	36.95
BG(1)-RP[d]	6.90	0.36	0.64	0.48
BG(52)-Rp[d]	13.52	2.71	1.91	2. 91
BG(1)-WP[d]	1.22	2.76	2.81	1.38
BG(52)-WP[d]	11.42	3.11	2.49	3.45

AIC is the Akaike Information Criteria
BIC is the Bayesian Information Criteria
SupLM is the Statistic to test for asymmetry
BG(i) is the Breush-Godfrey test for autocorrelation of order i (Critical value at the 5% level of significance is 3.84)

In the second subsystem (WP-RP, equation 19) the feedback effect is observed in both regimes. On the other hand, in the first regime the retail price reacts quicker to changes in the long-run equilibrium, while the opposite takes place during the second regime. In any event, these results can be better observed and understood by computing the impulse response functions.

3.3. Short-Run Dynamics

Short-run dynamic has 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 NIRF have been calculated for each subsystem mentioned above. 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, could expect different responses depending on which of the two regimes the shock is produced. If the cumulative response to positive and negative shocks is different from zero in absolute values, then we can conclude that the adjustment processes are asymmetric.

In order to analyse the asymmetric behaviour of price adjustments, the NIRF have been compute for $\delta=\pm0.5$, ±1 and $\pm$ 2 and for history-specific regimes such that the long-run equilibrium relationship [$\omega_t(\beta)=\beta' P_{t-1}$] is above or below the threshold value. Figures 1 and 2 show the NIRF for each system. 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 and for each asymmetry coefficient (equation 12), we have computed the corresponding 95% confidence intervals using bootstrapping techniques based on 5,000 replications.[12]

Figure 1 shows the NIRF for the subsystem (FP-WP) to a 1% positive and negative shock produced in both the first (Panel a) and the second (Panel b) regime[13]. Several implications for price relationships arise from it. As can be observed, the responses are fairly symmetric. Furthermore, The effects of positive and negative shocks are of the same magnitude. The asymmetric responses calculated as in (12), and considering different shock sizes, were non-significant and, thus, are not presented here. This symmetric behaviour is quite consistent with previous expectations, given that there is a direct link between farm and wholesale markets in Spain. This is the reason why both types of markets are taken in order to obtain a representative market price for Spain to calculate the ewe premium.

Positive (negative) shocks generate positive (negative) responses as expected in vertically related markets. In the long run we can note that the responses are non-significant, indicating that, as mentioned in the previous section, prices are homogeneous. In all cases, shocks generate permanent adjustments which are mostly complete after 16 weeks. The effect of a shock to the wholesale price on the farm price is non-significant during the first three weeks under both regimes, and thereafter gradually increases. However, when we consider a shock to the farm price, the wholesale price reacts quickly and significantly. These results suggest that, in the very short-run wholesale markets are more flexible to demand shocks. This can be partially explained by the existence of some contracts for future delivery, which make it impossible for farmers to react in the short-run to changing demand conditions. In any event, it seems that there is a strong interrelationship at these two levels of the marketing chain. The Results from the NIRF also seem

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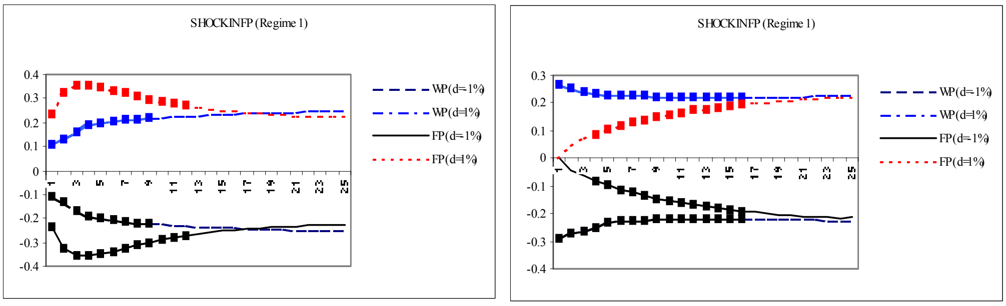
[12] All analyses have been carried out in GAUSS. We are grateful to Dr. van Dijk for providing valuable information on haw to tackle this cumbersome task.

[13] Significant responses are marked with a black square.

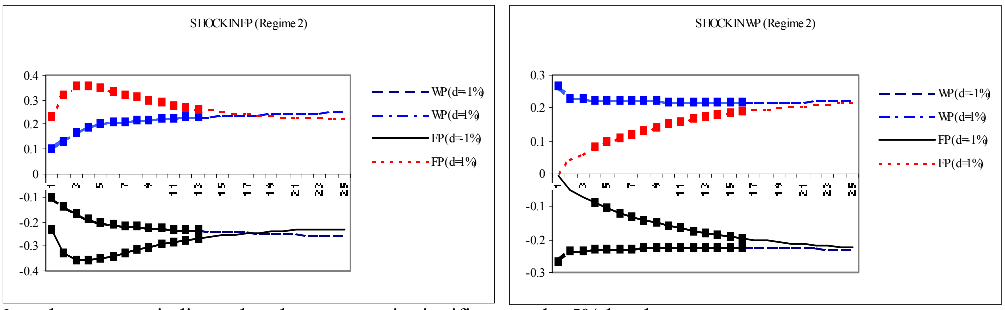
to reflect a cost-push transmission mechanism; that is, price transmission appears to occur mainly from farm to wholesale markets.

Figure 1. Non-linear generalised impulse responses function for the subsystem (FP-WP)

Panel a: Responses to positive and negative shocks in the first regime ($\omega_{t-1}(\hat{\beta}) \leq 0.057$)



Panel b: Responses to positive and negative shocks in the second regime ($\omega_{t-1}(\hat{\beta}) > 0.057$)



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

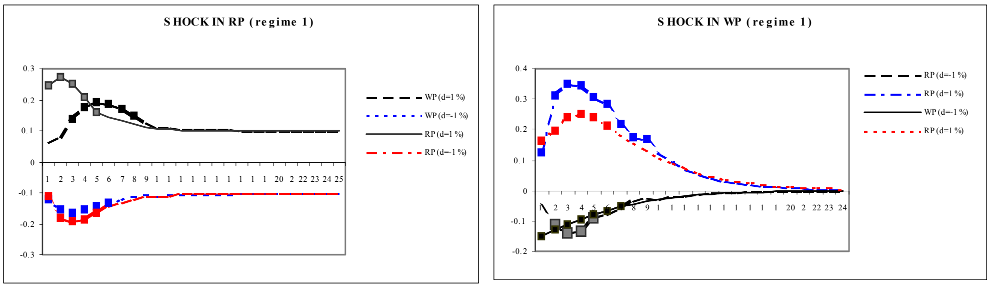
Let us now consider the subsystem (WP-RP) under the first regime, i.e. when prices are increasing (Figure 2, Panel a). A positive shock to the retail price does not generate any significant response on the path of the wholesale price in the first two weeks, while the response is significant when the shock is negative. On the other hand, if we observe the reaction of both prices to a positive shock to the retail price, although the reaction path is similar, the wholesale prices exhibit a certain delay in adjusting to the new situation. Thus, although in the long-run both prices are homogeneous, in the very short-run retailers benefit from a demand shock as the price spread increases by 50%.

A positive shock in the wholesale market generates an immediate response of the retail price of about 70% of the initial magnitude of the shock. Note, also, that although the time path of responses of both prices is similar, the magnitude of the response of the retail price is larger, indicating that retailers have the market power in the lamb market in Spain, as is the case with most perishable products. The distribution level is a highly concentrated sector, which allows retailers to react quickly, and, what it is more important, to overreact to supply shocks, increasing the price spread.

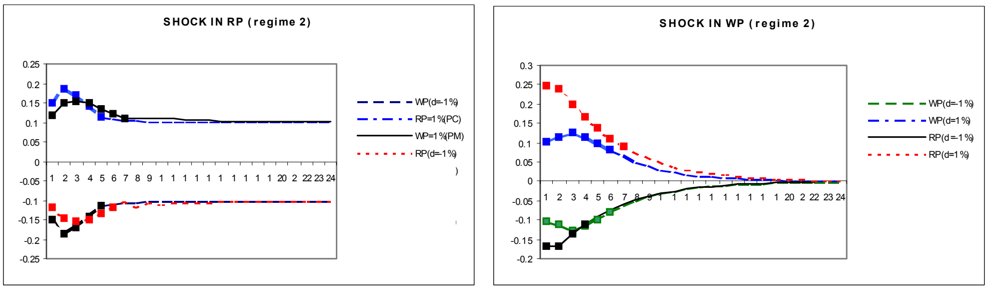
If we analyse the responses to negative shocks to the two prices in first regime, quite difference concerning the magnitude and the speed of the adjustment is appreciated. A negative shock in the retail price generates an immediate decrease in the wholesale price, while the retail price takes two weeks to react to a negative shock in the wholesale price. Both results indicate an increasing price spread in the very short-run, which benefit retailers. Finally, as can easily be observed, comparing the responses to positive and negative shocks, the price adjustment process is asymmetric at these two stages of the lamb marketing chain.

Figure 2. Non-linear generalised impulse responses function for the subsystem (RP-WP)

Panel a: Responses to positive and negative shocks in the first regime ($\omega_{t-1}(\hat{\beta}) \leq 0.044$



Panel b: Responses to positive and negative shock in the second regime ($\omega_{t-1}(\hat{\beta}) > 0.044$

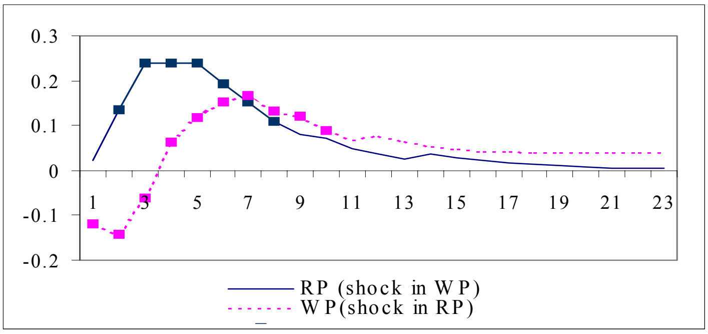


Note that squares indicate that the response is significant at the 5% level

The picture is somewhat different under the second regime, i.e. when prices are falling (Figure 2, Panel b). In this case, the convergence towards the long-run equilibrium takes place more quickly, especially when the shock is positive. Although the adjustment process is also asymmetric, as far as the price transmission mechanism is concerned, markets seem more integrated (the responses of the two prices to supply and demand shocks are of the same magnitude, mainly when the shock is positive).

Finally, and with the aim of summarising the results described above for the second subsystem, Figure 3 shows the asymmetric behaviour of the responses of retail and wholesale prices to a shock in any of both prices. In this case, the responses are computed as the sum of the NIRF for shocks of magnitudes ±0.5%, ± 1% and ±2%. This Figure might be useful to determine whether negative shocks are more persistent than positive shocks, or vice versa. Significant responses are marked with a black square. A shock in the wholesale price generates a positive asymmetric on the path of the retail price for all forecast horizons. This result suggests that retail prices show more nominal flexibility when they are increasing. However, in the case of the wholesale price, we can appreciate a negative asymmetric behaviour during the first three weeks, indicating a slow adjustment process when the retail price increases. Thereafter, the asymmetry is positive, reaching a maximum after seven weeks and being significant up to ten weeks (a little longer than in the case of the retail price). These results indicate that positive shocks are more persistent than negative shocks. 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 in turn give rise to increases consumers prices for food products and, therefore, to general price level.

Figure 3. Asymmetric coefficients of the system (RP-WP)



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

4. CONCLUSIONS

This paper has explored the non-linearity in the price transmission mechanism along the lamb marketing chain in Spain. Two subsystems have been studied: on one hand, the relationship between farm and wholesale prices; on the other, the price relationship between the wholesale and the retail market levels. The methodology used has been based on recent developments in non-linear adjustment models in a multivariate framework. The results point to a number of interesting conclusions.

In the long run, prices at different levels of the marketing chain are perfectly integrated; that is to say, any change in any of the prices is fully transmitted to the rest. This seems to be indicating that markets are fairly competitive, at least in the long run. However, in the short-run retailers benefit from any shock whether, positive or negative, that affects supply or demand conditions. This result is consistent with high degree of concentration that exists at the retail level in Spain.

Price adjustments between the farm and the retail level are symmetric and, moreover, they suggest that there is a cost-push transmission mechanism. Wholesalers react quickly to any change at the farm level, while farmers taking two weeks to react to changing demand conditions. In any case, there is a strong price relationship between the two market levels. The analysis of the price transmission mechanism between the wholesale and the retail levels offers more interesting results. In an environment of increasing prices, retailers are able to increase the price spread, independent of whether the supply or demand shocks are positive or negative. However, the situation changes substantially in the case of decreasing prices, where prices react more rapidly and the long-run equilibrium is achieved more quickly. Furthermore, asymmetries are not so evident.

The results presented in this paper are, of course, capable of improvements. First, we have not linked asymmetries with theory. New applications to the same sector in other countries with different market structures would allow linking our results with market power or holding stocks policy. Similarly, extending this application to other agricultural products with different degrees of vertical integration and/or horizontal concentration would enable us to obtain some conclusions about their effect on price transmission mechanisms. 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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