Price volatility and European food quality schemes

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TABLE OF CONTENTS

EXECUTIVE SUMMARY **4**

LIST OF TABLES **6**

LIST OF FIGURES **7**

LIST OF ABBREVIATIONS AND ACRONYMS **8**

1. INTRODUCTION **9**

2. METHODOLOGY **10**

2.1. Modelling nonstationary price time-series 10

2.2. Modelling time-varying volatility 14

3. THREE CASE STUDIES **15**

3.1. The Spanish beef sector 15

3.1.1. An overview of the PGI “Ternera de Navarra” 15

3.1.2. Data and Results 15

3.2. The Spanish lamb sector 23

3.2.1. An overview of the PGI “Cordero de Navarra” 24

3.2.2. Data and Results 24

3.3. The Italian cheese sector 31

3.3.1. An overview of the PDO “Parmigiano Reggiano” 31

3.3.2. Data and Results 32

4. CONCLUDING REMARKS **39**

REFERENCES **42**

EXECUTIVE SUMMARY

Food price volatility became more the norm than the exception since the 2008 financial crisis. It is common that prices of agricultural and food products fluctuate as market conditions change. Even more, price fluctuations can be considered as a signal of good performance of food markets. But, when these fluctuations are large and unanticipated, actors in the food marketing chain are challenged since adopting long-term decisions become more risky, generating a negative impact on the food security of farmers and consumers.

Since the 1980s, the European Union (EU) has regulated the quality schemes for agricultural products and food with the objective of helping producers to communicate to buyers and consumers the specific characteristics of such products and farming attributes, giving the possibility to producers to offer a unique and differentiated product of higher quality, normally at a higher price. However, most of the literature focuses on price premiums that consumers pay or are willing to pay. Moreover, no study has analysed the value generated along each stage of the food chain and to what extent farmers benefit from such higher consumer prices. And, more importantly, no study has dealt with the issue of to what extent food quality schemes (FQS) have contributed to reducing price volatility along the food chain. Probably the main reason of this lack of empirical literature is the lack of available data to conduct such type of analysis related to FQS products. In fact, we have not found any data source that provides systematic and continuous price information, both for conventional and FQS products, at the two extreme stages of the food supply chain, the producer and the retail level. For this reason, we believe that efforts should be directed to improving the monitoring and collection of this sort of price data.

The main aim of Task 4.3 of the STREGTH2FOOD project was to assess how prices for different FQS products are transmitted along the food marketing chain and the extent to which FQS have proved to be useful in reducing price volatility. Deliverable 4.4 presents the main results that were obtained. The methodology is based on the cointegration notion and after ensuring time-varying volatility in the data, a multivariate generalized autoregressive conditional heteroscedasticiy (MGARCH) is estimated. Our approach allows for possible asymmetry in the variance matrix.

The three case studies analyse FQS products in two Mediterranean countries, Spain and Italy. For Spain, we examine two Protected Geographical Indication (PGI) products, “Ternera de Navarra” (beef from Navarra) and “Cordero de Navarra” (lamb from Navarra). For Italy, we analyse the Protected Designation of Origin (PDO) “Parmigiano Reggiano”.

The deliverable draws conclusions on whether European FQS have proven to be useful in reducing price volatility along the food marketing chain. The results corroborate the presence of long-run relationships in the three case studies, with faster responses to deviations from the equilibrium in the short-run in the Spanish FQS than in conventional systems, whereas responses are faster in the Italian conventional case. Moreover, the results corroborate the existence of time-varying volatility with asymmetries. In general, the magnitude of price volatility patterns is higher in conventional systems than in FQS, and in retail markets (for beef and cheese products). Finally, results indicate that, at least for these three case studies, asymmetric dynamics are more significant in the conventional system, which favours the role of European FQS in reducing price volatility linkages between chain actors.

From the perspective of agriculture and food policy agents, the interest of our results is unquestionable because it improves knowledge about European FQS. Also, the difficulties in finding sources of reliable data, supports the need for further resources and efforts to monitoring, collecting and composing a reliable database of price series for premium products

and respective conventional counterparts, at least, at two different stages of their respective supply chains. The lack of available data in official databases prevent researchers from performing this kind of pair-wise analysis to gain a better understanding of the differences between product types.

LIST OF TABLES

Table 1. Summary of descriptive statistics for the Spanish beef price series 17
Table 2. Results of the integration order for the Spanish beef price series............................ 17
Table 3. VAR lag-length selection for the Spanish beef sector.. 18
Table 4. Results of the Co-integration analysis for the Spanish beef sector.......................... 18
Table 5. Results for the conditional mean model for the Spanish beef sector 19
Table 6. Results for the conditional variance model for the Spanish beef sector.................... 20
Table 7. Conditional variance equations for the Spanish beef sector.................................... 23
Table 8. Summary of descriptive statistics for the Spanish lamb price series 25
Table 9. Results of the integration order for the Spanish lamb price series............................ 26
Table 10. VAR lag-length selection for the Spanish lamb sector 27
Table 11. Results of the Co-integration analysis for the Spanish lamb sector........................ 27
Table 12. Results for the conditional mean model for the Spanish lamb sector 28
Table 13. Results for the conditional variance model for the Spanish lamb sector 29
Table 14. Conditional variance equations for the Spanish lamb sector 31
Table 15. Summary of descriptive statistics for Italian cheese price series 33
Table 16. Results of the integration order for the Italian cheese price series 33
Table 17. VAR lag-length selection for the Italian cheese sector.. 34
Table 18. Results of the Co-integration analysis for the Italian cheese sector........................ 35
Table 19. Results for the conditional mean model for the Italian cheese sector..................... 36
Table 20. Results for the conditional variance model for the Italian cheese sector 36
Table 21. Conditional variance equations for the Italian cheese sector 38

LIST OF FIGURES

Figure 1. Evolution of Spanish beef sector price series 16
Figure 2. Predicted volatilities and conditional correlations for the Spanish beef sector 22
Figure 3. Evolution of Spanish lamb sector price series 25
Figure 4. Predicted volatilities and conditional correlations for the Spanish lamb sector 30
Figure 5. Evolution of the Italian cheese sector price series 32
Figure 6. Predicted volatilities and conditional correlations for the Italian cheese sector 38

LIST OF ABBREVIATIONS AND ACRONYMS

AIC - Akaike's information criteria

ARCH - Autoregressive Conditional Heteroscedasticity

BEKK - Babba, Engle, Kraft and Kroner parameterization

BFGS - Broyden, Fletcher, Goldfarb and Shanno algorithm

BIC - Schwarz's information criteria

CC - Conditional Correlation

DF - (Augmented) Unit root test by Dickey and Fuller (1979)

ECT - Error Correction Term

EU - European Union

FP - Farm Price of the conventional food product

FPI - Farm Price of the food product designated with PGI label

GARCH - Generalized Autoregressive Conditional Heteroscedasticity

HQ - Hannan-Quinn's information criteria

IC - Information Criteria

KPSS - Stationarity test proposed by Kwiatkowski et al (1992)

MAIC - Modified Akaike's Information Criteria

MIC – Modified Information Criteria

MSB - Modified Sargan Bhargava unit root test

PDO - Protected Designation of Origin

PGI - Protected Geographical Indication

RC - Regulatory Council

RP - Retail Price of the conventional food product

RPI - Retail Price of the food product designated with PGI label

TSG - Traditional Speciality Guaranteed

VAR - Vector Autoregressive Model

VECM - Vector Error Correction Model

On Price Volatility and European Food Quality Schemes

H. Ferrer-Pérez and J.M. Gil [1]

1. INTRODUCTION

After the 2008 financial crisis, food price volatility became more the norm than the exception. In agricultural and food products, it is common that prices fluctuate due to changing market conditions. Even more, price fluctuations can be considered as a signal of good performance of food markets. However, when these fluctuations are large and unexpected all actors in the food marketing chain are challenged as adopting long-run decisions becomes more risky, generating a negative impact on the food security of farmers and consumers.

Since the 1980s the EU has regulated the quality schemes for agricultural products and food, seeking to help producers to communicate to buyers and consumers the specific characteristics of such products and farming attributes, protecting them from inferior copycat versions. In other words, producers are able to offer a unique and differentiated product of higher quality and, normally, higher price. Deselnicu et al. (2013) explore the main factors affecting the price premium associated with Food Quality Schemes (FQS). However, most of the literature focuses on price premiums consumers pay or are willing to pay (Aprile et al, 2012). And, more importantly, no study has dealt with the issue of to what extent FQS have contributed to reducing price volatility along the food chain. This is precisely the main aim of Task 4.3 of the STREGTH2FOOD project from which this Deliverable 4.4 shows the main results.

As the number of farmers producing the FQS product is significantly lower than in the case of the conventional counterpart and retailers are not able to buy these products in geographically separated markets, this is expected to result in a reduction in the market power at the retail level; an issue that has been treated extensively in the literature. As a consequence, we expect price fluctuations, due to unexpected supply and demand changing conditions, of the same magnitude both at farm and retail levels. Moreover, we expect a quick volatility transmission along the food supply chain in the case of FQS products.

In spite of the interest in this topic, there has not been any attempt in the past to address this issue. Assefa et al (2015) made a literature review of studies dealing with price volatility transmission in vertically related markets. From this review, we conclude that the literature is scarce; and second, no empirical studies exist dealing with FQS products. Probably the main reason is the lack of available data to conduct such analysis for FQS products. In fact, we have not found any data source that provides systematic and continuous information at the two extremes of the food supply chain. For this reason, we believe that efforts should be directed to improve monitoring and collection of this sort of data.

In Spain we have only found a reliable database for two PGI products: "Ternera de Navarra" and "Cordero de Navarra". Data contains farm and retail prices for the protected and conventional counterpart covering a six-year period. Data ranges from 2011 to 2015, and contains prices at farm and retail levels for both the FQS and conventional products. The methodological framework used is based on the Multivariate Generalized Autoregressive Heteroskedasticity (MGARCH) specification, in which asymmetries are allowed.

This deliverable is structured in three additional sections. Section 2 describes in more detail the methodological framework used, which is common to all case studies. Section 3 shows the main

[1] The authors would like to thank Fadi Abdelradi for insightful comments.

results obtained in each of the three case studies. This deliverable ends with some concluding remarks.

2. Methodology

Since agricultural products are usually characterized by a high time-varying volatility and a common trend over time, we conduct a detailed time-series econometric analysis to commodity prices so that we can assert the existence of co-movement in each pair of prices for the two aforementioned systems prior to modelling and testing for the possible presence of volatility spill-over effects in each market.

To do so, our approach consists of two parts: first, we carry out a thoughtful univariate analysis of the stochastic properties of the price series to determine the order of integration; and, second, once, we have determined the stochastic properties of the price series, we analyse the price volatility transmission in the short-run. This part involves co-integration analysis based on the maximum likelihood approach of Johansen (1988) and the analysis of the volatility spill-overs themselves. Note that a failure in the specification of the co-integration model will lead to spurious results which would invalidate the rest of the analysis.

2.1. Modelling nonstationary price time-series

Generally, the order of integration of time-series is analysed with the use of unit root tests. Not until the 1980s, did economists believe that economic series could be characterized as trend stationary, that is, series moving around a deterministic trend. However, this result was criticized in the influential work of Nelson and Plosser (1982) who proved that most macroeconomic time-series analysed in their paper should be treated as non-stationary in mean. Since then, a vast number of papers focus on the analysis of nonstationary variables. Diebold (1999), Engle and Granger (1987), Hamilton (1994), Maddala and Kim (1998), Phillips and Xiao (1998), Hayashi (2000) and Choi (2015) are, among others, excellent references.

Testing for unit roots in time-series has become necessary to establish links among prices because the test statistics behave differently depending on the stationary or non-stationary nature of the variable.

Surprisingly, a detailed review of the most recent empirical studies dealing with price transmission and price volatility of agricultural commodities shows that the implementation of the (augmented) test of Dickey and Fuller (1979), DF hereafter, and the PP tests of Phillips and Perron (1988) are greatly favoured at the expense of the tests proposed by Ng and Perron (2001) as a modified version of the DF and PP tests, based on the results obtained in Elliott et al (1996). This fact is quite surprising because the unit root literature, demonstrates that the tests proposed by Ng and Perron (2001) outperform DF and PP tests - see for instance Haldrup and Jansson (2006) and Patterson (2011),

So, in this study, according to the most recent contributions to the unit root literature, we consider the tests proposed in the influential work of Ng and Perron (2001). To justify our choice, we briefly explain their contribution below. Ng and Perron (2001) attempt to resolve two issues widely discussed in the literature. The first issue concerns the low power of standard unit root tests like the test of Dickey and Fuller (1979), and Phillips and Perron (1988) when the root of the autoregressive polynomial is close to unity. The second issue concerns the size distortions of most of the standard unit root tests when the moving-average polynomial of the first difference of the series has a large negative root. Their objective is twofold. First, they enhance the power of the tests less affected by the size-distortion, the M-tests proposed by Stock

(1999), using local GLS-de-trended (de-meaned) data as in Elliott et al (1996)[2]. Second, they derive a modified lag length criterion to determine the truncation lag parameter of the augmented Dickey-Fuller regression required to construct the autoregressive long-run variance estimator defined as:

$$\hat{\omega}_{AR}^2 = \hat{\sigma}_\varepsilon^2 \left(1 - \hat{\phi}(1)\right)^{-2} \tag{1}$$

where $\hat{\sigma}_\varepsilon^2 = T^{-1} \sum_{t=k+1}^{T} \hat{\varepsilon}_{tk}$ and $\hat{\phi}(1) = \sum_{i=1}^{k} \hat{\phi}_i$ with $\hat{\phi}_i$ and $\hat{\varepsilon}_{tk}$ obtained from the OLS augmented Dickey-Fuller regression:

$$\Delta\hat{y}_t = (\rho - 1)\hat{y}_{t-1} + \sum_{i=1}^{k} \phi_i \, \Delta\hat{y}_{t-i} + \varepsilon_t \tag{2}$$

where $\varepsilon_t \sim iidN(0, \sigma_\varepsilon^2)$ and $\hat{y}_t$ represents the generic filtered series. They show that the standard lag order selection methods like AIC and BIC underestimate the cost of selecting a small number of lags when the root of the moving-average polynomial is large and negative. Then, Ng and Perron propose a new class of modified information criteria which depend on a penalization factor which varies with the sample, the Modified Information Criteria (MIC) defined as:

$$MIC(k) \coloneqq \ln(\hat{\sigma}_k^2) + \frac{C_T(\tau_T(k) + k)}{T - k_{max}} \tag{3}$$

with $\tau_T(k) = (\hat{\sigma}_k^2)^{-1} \hat{\phi}_0^2 \sum_{t=k_{max}+1}^{T} \hat{y}_{t-1}^2$. If $C_T = 2$, we obtain the MAIC criterion and the MBIC is obtained with $C_T = \ln T$. Note as wel that both must satisfy $C_T/T \to 0$ as the sample size gets larger.

In our studies we apply the Modified Sargan Bhargava test, MSB henceforth, firstly proposed in Stock (1999) and later improved in Ng and Perron (2001). Our choice is based on its simplicity and remarkable size-power trade-off. The test is defined as follows:

$$\mathrm{MSB} = \frac{T^{-2} \sum_{t=2}^{T} \hat{y}_t^2}{\hat{\omega}_{AR}^2} \tag{4}$$

However, the outcome of the unit root tests developed in Ng and Perron (2001) is not valid when a structural break exists in the observed series as they are biased towards the non-rejection of the null hypothesis. To overcome this issue, it would be advisable to test for unit roots allowing for the existence of a single or multiple structural breaks.

Although, multiple alternatives are found in the unit root literature to address this issue[3] here we use the procedure developed in Carrion-i-Silvestre et al (2009) to test for the presence of

[2] If the local-to-unity alternative hypothesis is defined as $\rho_c = 1 + cT^{-1}$where c reflects the noncentrality parameter following Elliott et al (1996), these authors define the local-GLS procedure to detrend (demean) the series $\{y_t\}$ and the unknown deterministic vector $z_t$as follows: $y^{\bar{c}} = (y_1, (1 - \rho_{\bar{c}}L)y)'$ and $Z^{\bar{c}} = (z_1, (1 - \rho_{\bar{c}}L)z_{t-1})'$, being $\rho_{\bar{c}} = 1 + \bar{c}T^{-1}, \bar{c} < 0$ and L the lag operator so that $Lx_t = x_{t-1}$ for any given series x_t. Note that $\bar{c}$ is chosen so that the asymptotic local power function of the unit root test is tangent to the power envelope at 50% power.

[3] See Perron (2017) for an interesting editorial on the topic.

multiple unknown structural breaks in the level, intercept or intercept and slope of the series. In this study, we allow for two structural breaks at most.

As the traditional hypothesis testing favours the null hypothesis, we now consider the opposite set-up. We test the null of stationarity against the alternative of the existence of a unit root in the series. We apply the popular KPSS stationarity test developed in Kwiatkowski et al (1992).

As argued before, many commodity prices are usually characterized as co-integrated (Myers, 1994), implying that non-stationary prices share a trend in the long-run, and there are no incentives to deviate from this situation. But, if any unanticipated shock appears, there is a tendency to revert to the equilibrium.

Two widely used approaches are available in the related literature. On the one hand, the approach developed by Engle and Granger (1987) which relies on a two-step estimator to test the parameters of a bivariate single-equation model; and, on the other hand, the Johansen (1988) approach, which consists of a maximum likelihood ratio test to test multiple co-integrating vectors. Here, we follow the latter.

The procedure developed by Johansen (1988) is based on the link between the rank of a matrix and its characteristic roots[4]. So, the starting point is the correct specification of a vector autoregressive model VAR(k) with k denoting the optimal number of lags[5]. In this model, the variables are treated as endogenous and symmetrical (Sims, 1980). We write then[6]:

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + \cdots + A_k Y_{t-k} + \varepsilon_t \tag{5}$$

being $Y_t = \left(Y_{1t}, Y_{2t}, \dots, Y_{pt}\right)'$ a $p \times 1$ vector of endogenous variables where p is the number of variables; A_i for $i = 1,2,\dots,k$ are $(p \times p)$ matrices of autoregressive parameters. Also, ε_t is the error term in array form with $E(\varepsilon_t) = 0, \forall t$ and $E(\varepsilon_t \varepsilon_s) = 0$ for $t \neq s$; H for $t = s$, where H is the $(p \times p)$ variance-covariance matrix, which is positive definite.

To select the optimal truncation lag parameter, k, in the VAR(k) model, we consider information criteria as they are normally utilized for model selection (Aznar, 1989).

It is useful to rewrite equation (5) in the form of a vector error correction model (VECM) as follows:

$$\Delta Y_t = \Pi Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + \cdots + \Gamma_{k-1} Y_{t-k} + \varepsilon_t \tag{6}$$

with $\Gamma_i = -\sum_{i=j+1}^{k} A_i$ for $i = 1, \dots, k-1$ and $\Pi = -\left(I - \sum_{i=1}^{k} A_i\right)$ where I is the $(p \times p)$ identity matrix.

Within this context, testing for co-integration between variables implies selecting the rank r of matrix Π. Thus, once we have correctly specified the deterministic components in the model (Juselius, 2006) we can arrive at three possible scenarios:

[4] One might interprete this intuitively as a multivariate generalization of the Dickey-Fuller test.
[5] The lag order choice in VAR models is normally carried out using the Akaike Information Criterion (AIC), the Schwarz Information Criterion (BIC) and/or the Hannan-Quinn Information Criterion (HQ).
[6] To ease discussion, the deterministic terms are not included.

- If $r = 0$, there is no co-integration, the matrix is null and we have the usual VAR model in first differences.
- If $r = p$, Y_t is stationary and, hence, applying OLS to (6) will be efficient.
- If $1 \leq r < p$, r denotes the number of co-integration relations with $\Pi Y_{t-1} \sim I(0)$ and so, we can decompose Π matrix as follows:

$$\Pi = \alpha\beta' \tag{7}$$

with β the matrix of parameters from the r co-integration relationships, and α measuring the speed of adjustment of the parameters towards the equilibrium in the long-run ($\beta' Y_{t-1}$).

Hence, selecting the co-integration rank is equal to determining the number of the characteristic roots of the matrix Π that differ from zero[7]. However, in practice, we can only estimate Π and its respective characteristic roots. Two following test statistics are available in the literature and defined as:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{p} \ln(1 - \hat{\lambda}_i) \tag{8}$$

$$\lambda_{max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1}) \tag{9}$$

where $\hat{\lambda}_i$ represents the estimated eigenvalues, that is, the values of the characteristic roots obtained from $\widehat{\Pi}$, and the sample size is indicated with T. Curiously, the null hypothesis tested differs for (8) and (9). S, we have for the λ_{trace} statistic:

$$\mathrm{H}_0: r_0 \leq r \qquad H_1: r_0 > r \tag{10}$$

and for the λ_{max} statistic[8]:

$$\mathrm{H}_0: r = r_0 \qquad H_1: r = r_0 + 1 \tag{11}$$

However, in this case, we employed the Bartlett corrected trace test λ^{*}_{trace} because we can ensure reduced size distortions in the trace tests due to the short-run effects of the VAR model (Johansen, 2002).

Once the long-run relations are found, the following natural step is to analyse whether this relationship is stable over the period analysed. In this sense, we can use the proposals of Hansen (1992) and Gregory and Hansen (1996a, 1996b), among others.

Now, we can estimate the VECM expressed in (6) adapted to our case as:

$$\Delta P_t = \mu + \alpha P_{t-1} \Pi P_{t-1} + \sum_{i=1}^{k} \Gamma_i \, \Delta P_{t-i} + \varepsilon_t \tag{12}$$

[7] If the variables in Y_t are not cointegrated, the rank is zero and subsequently all of the characteristic roots are zero.
[8] Critical values for both test statistics were computed using Monte Carlo simulations. For further details, the interested reader may consult Johansen (1988), Juselius (2006) and Enders (2010), among others.

where $P_t = (P_{1t}, P_{2t})'$ denotes the prices at the farm and retail levels of the supply chain for each system and the error term $\varepsilon_t|\Omega_{t-1} \sim iidN(0, H_t)$ with 2x2 variance-covariance matrix, H_t. Also, ΠP_{t-1}measures the long-run relation between commodity prices and Γ_i measures the short-run effects of the mean model.

Notwithstanding, the assumption underlying the estimation procedure above is that the variance of ε_t in (12) is constant over time, which may be quite unrealistic given that it is common to see periods in which the fluctuations may be unusually high followed by relative tranquillity.

2.2. Modelling time-varying volatility

As claimed before, price time-series usually exhibit time-varying volatility in the short-run, a characteristic which invalidates the conventional assumption of homoscedasticity. For instance, agricultural commodity prices can be seen as an obvious candidate.

Thus, agricultural economists may often think of multiple situations in which they are interested in measuring this volatility[9]. To do so, a widely used econometric technique is the Generalized Autoregressive Conditional Heteroscedastic (GARCH) model proposed by Bollerslev (1986), an extension of the seminal work of Engle (1982)[10] which allows the variance-covariance matrix to depend not only on lagged residuals but also on its own lags. Moreover, the literature provides several model specifications[11], but in this work we apply the Babba, Engle, Kraft and Kroner (BEKK) parameterization developed by Engle and Kroner (1995)[12] so that we can identify volatility spillovers across interrelated markets. Also, we follow Kroner and Ng (1998) to allow for asymmetric volatility patterns so as to ensure a correct specification of our models, because the original BEKK parameterization is restricted to symmetric effects. Therefore, our approach allows us to capture the possibility that volatility responses tend to be greater for negative shocks than for positive shocks.

The BEKK specification is as follows:

$$H_t = CC' + A'u_{t-1}u'_{t-1}A + B'H_{t-1}B \tag{13}$$

being C a lower triangular matrix of constants; A is a 2x2 matrix of ARCH term coefficients, and B is a 2x2 matrix of GARCH term coefficients. Subsequently, the extension developed by Kroner and Ng (1998) to capture the asymmetric effects is as follows:

$$H_t = CC' + A'u_{t-1}u'_{t-1}A + B'H_{t-1}B + D'v_{t-1}v'_{t-1}D \tag{14}$$

where D is a 2x2 matrix that measures the asymmetries, defining $v_t = u_t$ if u_t is negative and $v_t = 0$ if otherwise. Note that matrix A is a coefficient matrix for own and cross recent shock transmission effects, while the matrix B contains coefficients for own and cross past volatility transmission effects.

[9] In this study, we use indistinticly the concepts of volatility and variance. However, it should be noted that here we are interested in conditional variance which represents the short-run variance and takes the past information available to forecast.
[10] Engle (1982) was the first to simultaneously model the mean and the variance of a time-series.
[11] See an interesting review of the specifications employed in price volatility transmission in Assefa et al (2015).
[12] This approach ensure the positive definiteness of the covariance matrix by constraining these matrices to be symmetric.

Therefore, our strategy permits us to examine price volatility behaviour along the food supply chain allowing for asymmetric effects, which may be useful not only for agents directly and indirectly affected in markets but also for those involved in the design of food policies.

3. THREE CASE STUDIES

3.1. The Spanish beef sector

3.1.1. An overview of the PGI "Ternera de Navarra"

(La Comunidad Foral de) Navarra is located in the West Pyrenees leaning towards the river Ebro sharing a border in the north with France, in the south with La Rioja and Zaragoza (Spain); in the east with Zaragoza and Huesca (Aragón) and with Álava and Guipúzcoa (País Vasco), in the west. The area spans in total, 10.506 km, mostly mountainous terrain. The part designated for breeding the beef from Navarra represents almost 96% of the total territory. This geography together with the weather and agricultural characteristics makes Navarra a region of contrasts which favours the development of the Beef of Navarra.

PGI "Ternera de Navarra" (beef from Navarra)[13] was created in 1994 as a label of origin with the objective of protecting and promoting an underestimated foodstuff produced traditionally so that the consumer may perceive this system as ensuring the superior quality of the beef meat. The method of production is based on the sustainable exploitation of natural resources and the environment along with exhaustive controls realized by ENAC (national body that certifies the reliability of these controls). In 2000 Ternera de Navarra was designated with the European PGI label.

In 2016, the Regulatory Council, RC hereafter, registered 513 farms, 15 more than in the year 2015, and with 180 butchers authorized for commercialization. Farms raise the Pirenaica, Blonde, Parda Alpina, Charolais breeds and their crossbreeds too. Suckling is compulsory at least during the first four months after birth, where the suckler cow is allowed to graze as usual, and be fed a supplementary nutrition composed by maize, barley, soya, wheat, beans, and concentrated foodstuff authorized by the RC. The RC guarantees the quality of the PGI Beef of Navarra by controlling the maturing period of the beef (at least one week) until gaining the optimal conditions of colour, taste, aroma and tenderness with a pH always less or equal than 5.8, and by controlling the traditional cutting process. The RC has 4 slaughterhouses which slaughter 6.856 animals, producing around 2,177.246 tonnes of beef, mostly commercialized in the domestic market with small export figures to international markets.

3.1.2. Data and Results

Data employed to carry out the empirical study was collected from the Observatory of Agricultural Prices of the Government of Navarra. Data are available for the PGI "Ternera de Navarra" and its conventional counterpart so that we could compare both systems. Prices time-series are expressed in Euros, and available at farm level (price paid to the farmer) and retail level (price paid by the consumer). Our choice follows the natural selection commonly used in the price transmission literature (Goodwin and Piggott, 2001).

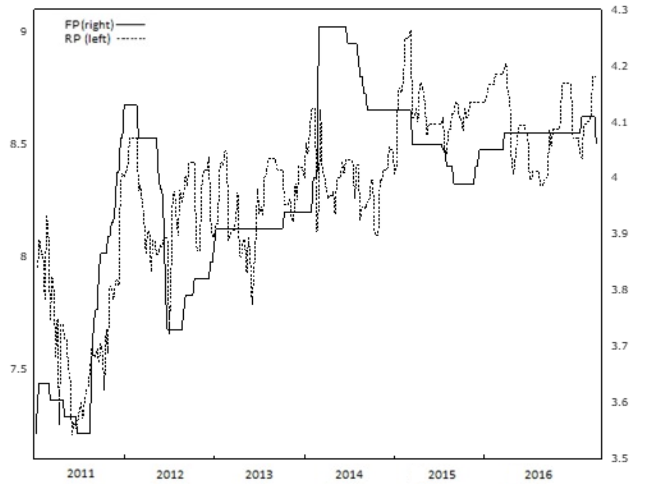
Weekly prices are available for the period from 2011 to 2016. For the quality system, we have farm (FPI) and retail (RPI) prices whereas for the conventional system, we have farm (FP) and retail (RP) prices. Therefore, our data set covers an important period after the recent rise in prices in 2007/2008. We present nominal prices in Figure 1, in which Panel A illustrates the conventional beef and Panel B illustrates the PGI beef. In Panel A, prices show a similar

[13] "Nafarroako Aratxea" in Euskera.

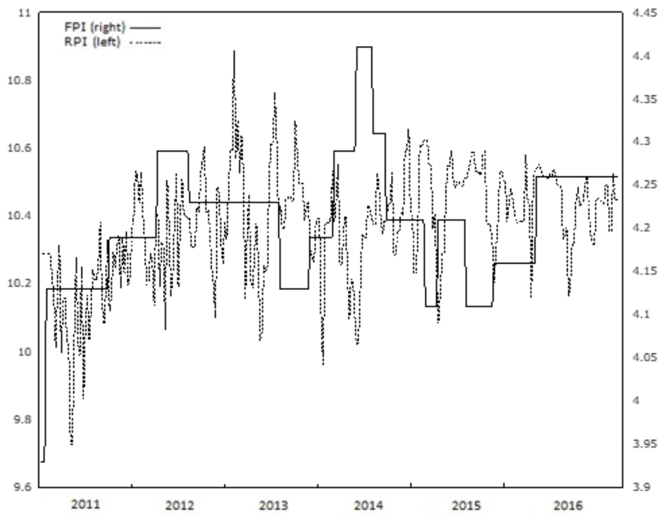
trending pattern during the period, which may suggest an equilibrium relationship in the long-run, but with more volatile periods at the retail level than at the farm level. In Panel B, prices also seem to obey a long-term relationship, again with more significant volatile episodes at retail than at farm prices. In this case, farm prices tend to adjust only a bit slower after a change at the retail level. Every price series exhibits visible volatility and a trending growth over the whole sample period. Moreover, all the price series were transformed into natural logs according to theory (Banerjee et al, 1993).

Figure 1. Evolution of Spanish beef sector price series

Panel A. Conventional "Ternera de Navarra"



Panel B. PGI "Ternera de Navarra"



Source: Own calculation based on Regional Government of Navarra (Spain), Observatory of Agricultural Prices database. Vertical axes are measured in €/kg carcass.

Table 1 shows a summary of the descriptive statistics of the series for each system. The statistics indicate that the trend is significant in both the conventional and quality scheme price series but almost negligible. All the prices are found to exhibit non-normality and ARCH effects. In all, we have 312 observations for each series. Unit root tests and stationarity tests confirm the existence of a unit root in the price series of the two systems, and this result is robust to the presence of structural breaks. Table 2 reports the results.

Table 1. Summary of descriptive statistics for the Spanish beef price series

	FP	RP	FPI	RPI
Mean	3.975	8.305	4.207	10.370
Median	4.050	8.379	4.210	10.385
Minimum	3.545	7.204	3.930	9.729
Maximum	4.270	9.010	4.410	10.888
Standard deviation	0.179	0.371	0.072	0.166
Skewness	-0.788***	-0.916***	-0.024***	-0.644***
Kurtosis (excess)	0.120	0.590**	1.880***	0.990***
Jarque-Bera test	32.487***	48.176***	45.941***	34.307***
Engle (1982)'s test	286.070***	198.099***	224.804***	62.333***
Trend	1.325e $^{-03}$***	3.159e $^{-03}$***	1.981e $^{-04}$***	8.231e $^{-04}$***
# observations	312	312	312	312

Notes: We have considered logarithmic transformations of the prices in our application. The use of the test of Engle (1982) allows us to check whether there are ARCH effects. In this case, we have used 2 lags. *** (**) denotes statistically significant at 1% (5%) level of significance.

According to theory, co-integration exists when nonstationary variables, prices in our case, show a tendency to move together in the long term and deviations from this equilibrium due to unanticipated shocks tend to revert eventually. Further, recall that the co-integration analysis is based on the unrestricted vector autoregressive (VAR) model. See Juselius (2006) for an excellent illustration of the co-integrated VAR model.

Before testing for co-integration, it is necessary to determine the number of lags to be included in the unrestricted VAR model. Table 3 reports the results.

Based on the lag choice for each system, we test for co-integration following Johansen (1988, 2002) and determine the co-integration rank[14] using the Bartlett corrected trace test λ^*_{trace}. Results, which are reported in Table 4, show the respective relationships in each system. As prices are considered in logs, we can read the co-integration parameters as price elasticities. In both cases, we can observe a positive relationship, which is especially strong in the conventional system (68.6%) implying that an increase in farm prices will lead to a rise in retail prices. Whereas for the quality system, farm prices exert very low influence on retail prices, about 7% of price elasticity.

Table 2. Results of the integration order for the Spanish beef price series

Panel A. Unit root and stationarity tests

	MSB	KPSS

[14] Note that if we apply the method suggested in Engle and Granger (1987) we also found the existence of respective co-integration relationships in the two systems.

FP	0.284 (3)	0.471 (1)**
RP	0.337 (3)	0.673 (0)**
FPI	0.284 (3)	0.471 (1)**
RPI	0.337 (3)	0.673 (0)**

Panel B. Unit root tests allowing for structural breaks (Carrion-i-Silvestre et al, 2009)

	ADF	MSB	ℓ	Tb
FP	-1.978 (-3.092)	0.300 (0.161)	3	2016:20 (281)
RP	-2.368 (-3.098)	0.201 (0.160)	5	2016:19 (280)
FPI	2.521 (-3.172)	0.398 (0.157)	3	2011:32 (32)
RPI	-3.097 (-3.827)	0.152 (0.130)	1	2012:44 (96) 2015:21 (229)

Notes: In Panel A, we apply the MSB unit root test as in Ng and Perron (2001) and the KPSS stationarity test. The truncation lag parameter, k, presented in parentheses and is estimated using the MAIC. Also, ** denotes statistically significant at 5% level of significance since the asymptotic critical values at 5% level for the trend case are respectively 0.168 (MSB) and 0.146 (KPSS). In Panel B, ℓ reflects the bandwidth parameter for the KPSS test selected with the automatic bandwidth procedure of Andrews (1991) for the kernel-based estimator of the long-run variance. The critical value at the 5% significance level of each test is shown in parentheses. Tb reflects the time breaks, that is, the date when a structural break was endogenously detected and the corresponding number of observation is presented in parentheses.

Table 3. VAR lag-length selection for the Spanish beef sector

Panel A. Conventional System

IC	**lags**
BIC	2
AIC	5
HQ	2

Panel B. Quality System

IC	**lags**
BIC	1
AIC	3
HQ	1

Notes: Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Table 4. Results of the Co-integration analysis for the Spanish beef sector

Panel A. Conventional System

Rank	Eigen value	λ^*_{trace}	p-value
0	0.039	16.211	0.037
1	0.013	3.833	0.050

Co-integration relationship: $ECT = RP_t - 1.170^{***} - 0.686^{***}FP_t$

Panel B. Quality System			
Rank	Eigen value	λ^*_{trace}	p-value
0	0.089	36.601	0.000
1	0.027	8.404	0.070

Co-integration relationship: $ECTI = RPI_t - 2.240^{***} - 0.068\,FPI_t$

Notes: *** denotes statistically significant at 1% level of significance. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

The results derived from the estimation of the VECM are shown in Table 5. As regards to the short-run dynamics of the model estimation, we can see that for the conventional system the α_i coefficients indicate that only the retail price responds to deviations from the long-run equilibrium by reverting to it at 8%, as also occurs in the quality system but at around 27%. Implying faster adjustments than in the conventional system. Moreover, conventional retail prices are affected by own first lagged values and farm price first lags. Whilst in the quality system, the retail price is only affected by own second lagged values. Farm prices in the conventional system are affected by retail and farm price first lags and the second own lag, but in the quality system we cannot find statistical significance of any lags.

We also ensure model adequacy in both systems by conducting a multivariate residual analysis. In the lower panel of Table 5 we report the results for the multivariate Q statistic of Hosking (1980, 1981) and the multivariate ARCH LM tests for each system. The former procedure tests the null of no autocorrelation of the residuals of the system via Portmanteau test, whereas the latter tests the null that the residual series have mean zero, and are not serially correlated with a fixed variance-covariance matrix via an LM type test. For the conventional system, the null of no autocorrelation of the multivariate Hosking's test cannot be rejected and we can assure the use of a multivariate GARCH specification. For the quality system, the diagnosis cannot ensure the presence of ARCH effects, maybe due to the construction of the test itself because some of the tested coefficients may be more likely to be informative than others. So, we should be a bit more careful with the use of only this test. Hence, to assess whether there are ARCH effects in each regression, we test for the presence of individual ARCH effects in the series. The results obtained indicate that we strongly reject the null of no ARCH effects in the retail equation but not in the farm equation, for which we have shown that no coefficients are statistically significant and are indeed biasing the result of the test and hence failing in identifying the real presence of ARCH effects in the residuals. The results indicate that there is time-varying volatility.

Table 5. Results for the conditional mean model for the Spanish beef sector

Short-run parameters for the Conventional System

$$\begin{pmatrix}\Delta RP_t\\ \Delta FP_t\end{pmatrix} = \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix} ECT_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121}\\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-1}\\ \Delta FP_{t-1}\end{pmatrix} + \begin{pmatrix}\delta_{112} & \delta_{122}\\ \delta_{212} & \delta_{222}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-2}\\ \Delta FP_{t-2}\end{pmatrix} + \varepsilon_t$$

Short-run parameters for the Quality System

$$\begin{pmatrix}\Delta \mathrm{RPI}_t \\ \Delta \mathrm{FPI}_t\end{pmatrix} = \begin{pmatrix}\alpha_1 \\ \alpha_2\end{pmatrix}\mathrm{ECTI}_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121} \\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta \mathrm{RPI}_{t-1} \\ \Delta \mathrm{FPI}_{t-1}\end{pmatrix} + \begin{pmatrix}\delta_{112} & \delta_{122} \\ \delta_{212} & \delta_{222}\end{pmatrix}\begin{pmatrix}\Delta \mathrm{RPI}_{t-2} \\ \Delta \mathrm{FPI}_{t-2}\end{pmatrix} + \begin{pmatrix}\delta_{113} & \delta_{123} \\ \delta_{213} & \delta_{223}\end{pmatrix}\begin{pmatrix}\Delta RPI_{t-3} \\ \Delta FPI_{t-3}\end{pmatrix} + \begin{pmatrix}\delta_{114} & \delta_{124} \\ \delta_{214} & \delta_{224}\end{pmatrix}\begin{pmatrix}\Delta RPI_{t-4} \\ \Delta FPI_{t-4}\end{pmatrix} + \varepsilon_t$$

	Conventional System		*Quality System*			
	$i = 1$	$i = 2$	$i = 1$	$i = 2$	$i = 3$	$i = 4$
α_i	-0.084*** (0.027)	0.003 (0.010)	-0.272*** (0.027)	-0.017 (0.018)		
δ_{11i}	-0.198*** (0.058)	-0.092 (0.057)	-0.115*** (0.030)	-0.059** (0.024)	-0.010 (0.043)	-0.057** (0.028)
δ_{12i}	0.374** (0.147)	-0.104 (0.146)	-0.104 (0.170)	0.013 (0.155)	-0.136 (0.136)	0.067 (0.093)
δ_{21i}	-0.063*** (0.023)	-0.013 (0.023)	0.018 (0.034)	0.028 (0.018)	-0.030* (0.018)	-0.028 (0.030)
δ_{22i}	0.199*** (0.058)	0.122** (0.058)	-0.012 (0.021)	0.005 (0.027)	-0.010 (0.045)	-0.008 (0.045)
Hosking's test: 35.460 (0.91)			Hosking (1980) test: 22.214 (0.99)			
Multivariate ARCH LM test: 69.000*** (0.00)			Multivariate ARCH LM test[a]: 6.65 (0.67)			

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for multivariate tests in which contains the p-value.
[a] ARCH LM test (lags=3) for the residuals of the retail equation (RPI on FPI) strongly rejects the null of no ARCH effect (p-value): 8.31** (0.04), whereas the residuals from the farm equation (FPI on RPI) show non-rejection of the null of no ARCH effect 1.64 (0.64), and hence, delivering a biased outcome of the test.

Table 6 reports the estimated results for the asymmetric multivariate BEKK-GARCH conditional variance model. Also, we show usual tests on standardized residuals that ensure a correct model specification. We checked that the eigenvalues were less than unity, which implies that the two estimated models are covariance stationary. Using the Nyblom (1989) test we ensure the stability of the volatility models for the conventional and quality systems, respectively.

For the conventional system, we checked the presence of time-varying volatility by rejecting the null hypothesis of parameters in matrices A, B, and D being equal to zero. We can claim that there are asymmetric effects as the null of parameters in matrix D equaling zero is strongly rejected, also for the quality system. We also check whether time-varying volatility is present in the quality system and we see that the LR test strongly rejects the null that parameters of matrices A and B are equal to zero. Joint stability can only be claimed at the 5% significance level with all the coefficients individually stable in the system, at least, at the 1% significance level. Summing up, both in the conventional and the quality scheme systems, the presence of time-varying volatility is corroborated by the data.

Table 6. Results for the conditional variance model for the Spanish beef sector

GARCH model parameters

Asymmetric BEKK for the Conventional System

$$C = \begin{pmatrix}c_{11} & 0 \\ c_{21} & c_{22}\end{pmatrix}, A = \begin{pmatrix}a_{11} & a_{12} \\ a_{21} & a_{22}\end{pmatrix}, B = \begin{pmatrix}b_{11} & b_{12} \\ b_{21} & b_{22}\end{pmatrix}\ and\ D = \begin{pmatrix}d_{11} & d_{12} \\ d_{21} & d_{22}\end{pmatrix}$$

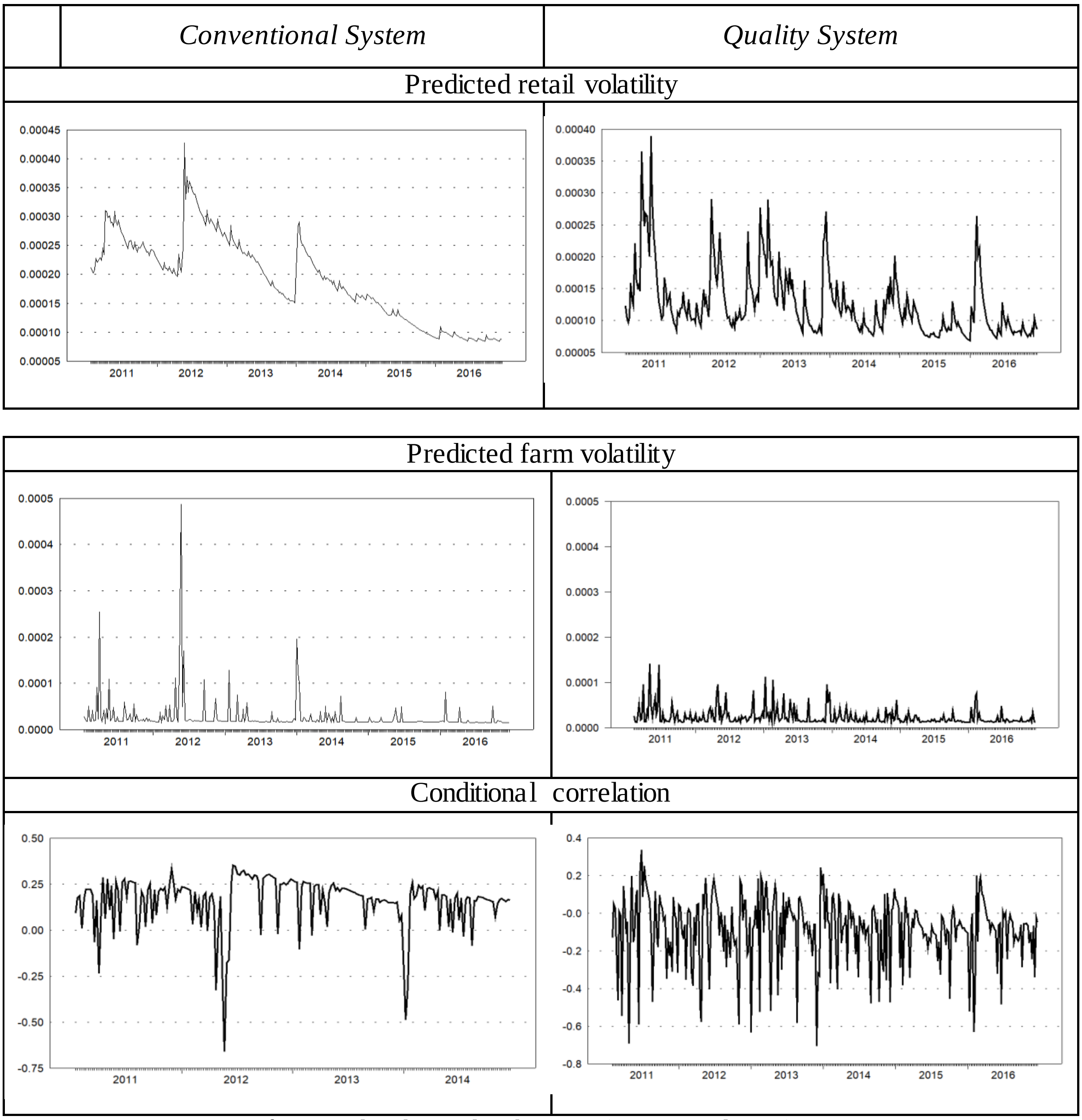
Conventional System	Quality System

	$i = 1$	$i = 2$	$i = 1$	$i = 2$
c_{1i}	0.002***(0.000)		0.003*** (0.000)	-
c_{2i}	-0.004*** (0.000)	0.000 (0.001)	-0.002*** (0.000)	0.003*** (0.000)
a_{1i}	0.111*** (0.020)	0.012 (0.030)	0.187*** (0.022)	-0.295*** (0.015)
a_{2i}	-0.156** (0.064)	0.166*** (0.045)	0.060 (0.280)	-0.106 (0.093)
b_{1i}	0.968*** (0.005)	0.086*** (0.010)	0.891*** (0.006)	0.046* (0.025)
b_{2i}	0.186*** (0.017)	0.128** (0.066)	-0.273*** (0.070)	-0.044 (0.106)
d_{1i}	0.089*** (0.040)	-0.246*** (0.068)	0.339*** (0.040)	-0.012 (0.110)
d_{2i}	-0.439*** (0.117)	0.689*** (0.218)	0.005 (0.085)	-0.003 (0.161)
LR test for the null of joint significance of parameters of matrices A,B,D: 86149.859*** (0.00)			18675.00*** (0.00)	
LR test for the null of joint significance of parameters of matrices A,B: 65776.313*** (0.00)			19805.29*** (0.00)	
LR test for the null of asymmetric effects:		8.932*** (0.00)	26.948*** (0.00)	
LR test for BEKK cross effects:		415.090*** (0.00)	156.90*** (0.00)	
Nyblom (1989) joint stability test:		3.065 (0.16)	7.24* (0.06)	
Stable roots of the system?		YES	YES	

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for the multivariate tests reported in the lower panel in which contains the p-value. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Figure 2 illustrates the predicted conditional variances for the conventional (left panel) and quality (right panel) systems, respectively. It can be seen that volatilities are not constant over time, especially in the conventional system in which the volatility is higher than in the quality system. This can be noticed by comparing the magnitudes of the respective predicted volatilities in the left- and right-hand side panels of Figure 2. In the left panel, we can observe a negative trend in the volatility pattern at retail level which is transmitted to the farm level indicating that the volatility in the market has been steadily reducing during the 2011-2016 period. Moreover, we can see that volatility at farm level is characterized by sharp rises in the first weeks of 2011, 2012 and 2014, which may be interpreted as volatility spillovers from the retail level. Finally, by looking at the lower panel we can observe a strong relationship between market shocks, slightly higher in the conventional system. In both systems, conditional correlations are low and change from positive to negative abruptly which may be interpreted as a result of poor transmission of price signals or unexpected responses to negative values. This may indicate limited volatility spilled from retail prices.

Figure 2. Predicted volatilities and conditional correlations for the Spanish beef sector



Source: Own source from results obtained with CATS (Dennis et al, 2006) in RATS 9.0.

As we cannot directly interpret the individual coefficients estimated in the multivariate GARCH parameterization for each of the two systems, we discuss the conditional variance equations presented in Table 4.

For the conventional system, in the upper panel, we can see that the volatility in the retail price is not only directly affected by its own past volatility (h_{11t-1}), past farm volatility (h_{22t-1}) but also indirectly through the covariance terms (h_{12t-1}). This indirect effect implies that the high correlation between retail and farm levels is relevant and will help to decrease price instability, guaranteeing stability when prices move in the same direction. But, we can also interpret this fact as there being some retail power that affects instability in farm markets. Results also suggest that past shocks to retail and farm markets have an asymmetric effect on retail price volatility whereas only past shocks to the retail market have an asymmetric effect on farm price volatility (h_{22t}). Moreover, the volatility in each of the two markets is affected by their respective own shocks (u_{1t-1}^2, u_{2t-1}^2).

For the quality system (lower panel), the results indicate that volatility in the retail price (h_{11t}) is only affected by its own past volatility (h_{11t-1}), and it is not influenced by past farm volatility. Also, it seems that the correlation between retail and farm levels has no impact on price volatility (h_{12t-1} is not statistically significant). However, volatility in the farm price (h_{22t}) is directly affected by past retail volatility. It seems that past shocks originated at retail markets do affect retail and farm volatilities (u_{1t-1}^2 is statistically significant) by increasing market volatility.

Table 7. Conditional variance equations for the Spanish beef sector

Conventional System

$$h_{11} = 1.8\,e^{-05^{***}} + 0.937^{***}h_{11t-1} + 0.035^{***}\mathrm{h}_{22t-1} + 0.360^{***}h_{12t-1} + 0.012^{***}\,u_{1t-1}^2 + 0.024\,u_{2t-1}^2 - 0.035^{**}u_{1t-1}u_{2t-1} + 0.008^{*}v_{1t-1}^2 + 0.193^{*}v_{2t-1}^2 - 0.078^{***}v_{1t-1}v_{2t-1}$$

$$h_{22} = 0.0000 + 7.382\,e^{-03^{***}}h_{11t-1} + 0.016\,\mathrm{h}_{22t-1} + 0.022^{*}h_{12t-1} + 1.36\,e^{-04}\,u_{1t-1}^2 + 0.0277^{*}u_{2t-1}^2 - 0.035^{***}\,u_{1t-1}u_{2t-1} + 0.060^{*}v_{1t-1}^2 + 0.475\,v_{2t-1}^2 - 0.339^{**}v_{1t-1}v_{2t-1}$$

Quality system

$$h_{11} = 1.2e^{-05^{***}} + 0.794^{***}h_{11t-1} + 0.074^{*}\,\mathrm{h}_{22t-1} - 0.486^{***}\,h_{12t-1} + 0.035^{***}\,u_{1t-1}^2 + 0.004\,u_{2t-1}^2 + 0.022\,u_{1t-1}u_{2t-1} + 0.115^{***}v_{1t-1}^2 + 2.6\,e^{-05}v_{2t-1}^2 + 0.003\,v_{1t-1}v_{2t-1}$$

$$h_{22} = 1.0e^{-05^{***}} + 0.002\,h_{11t-1} + 0.002\,\mathrm{h}_{22t-1} - 0.004\,h_{12t-1} + 0.087^{***}\,u_{1t-1}^2 + 0.011\,u_{2t-1}^2 + 0.062\,u_{1t-1}u_{2t-1} + 1.35\,e^{-04}v_{1t-1}^2 + 7.0\,e^{-06}v_{2t-1}^2 + 6.2\,e^{-05}\,v_{1t-1}v_{2t-1}$$

Notes: h_{11} retail price, h_{22} farm price variance. Estimated parameters of h_{ijt-1},$i,j = 1,2$ reflects direct and indirect volatility transmission between prices, whereas those of u_{it-1}^2 and $u_{it-1}u_{jt-1}$ indicates how price volatility is affected by markets shocks. ***, ** and * denotes statistical significance at the 1%, 5% and 10% level, respectively. "e" indicates scientific notation (with exponent).

Hence, our results show that price linkages between chain actors of these two beef supply chains in Spain are characterized by volatility spill-overs, slightly superior in the conventional system. Causality found in the data seems to indicate that to some extent, market power is exerted by retailers. The evidence provided for the quality system suggests smaller volatility patterns and subsequently, more stable prices.

3.2. The Spanish lamb sector

Following a similar structure, we analyse the Spanish lamb sector. After a brief introduction of the food quality scheme, we present and discuss the results derived from this study.

3.2.1. An overview of the PGI "Cordero de Navarra"

The PGI "Cordero de Navarra" (lamb from Navarra)[15] was designated with the EU PGI label in 2003 pursuing a similar objective as with the PGI beef. This quality designation only protects lambs from the Navarra and Lacha breeds. We can distinguish two types of lambs: the suckling lamb ("cordero lechal"), which is only fed with milk from the suckler lamb, and the light lamb ("cordero ternasco"), which is fed with milk at least until 45 days after the birth for those of the Navarra breed and 25-30 for those of the Lacha breed, after which they are fattened with white cereal straw and a concentrate made mainly from cereals, legumes, vitamins and minerals. Both types of lamb are raised following traditional methods linked to territory based on extensive-type or semi-extensive-type systems, in which the diet is based on grass, fodder and cereals so the sustainability is guaranteed. In this study we focus on the semi-extensive system. The RC has more than 200 families which make a living in raising PGI lambs, more than 50.000 lambs are certified with the PGI label and 89 butchers authorized for commercialization in the region of Navarra.

3.2.2. Data and Results

As in the previous case study, we have extracted our data set from the official statistics supplied by the Observatory of Agricultural Prices of the Government of Navarra (Spain) which includes a recent five-year period after the rise in prices that took place in 2007/2008 covering the period 2011-2016 with a total of 312 observations. Weekly prices, expressed in Euros, are available at the farm and retail levels for the PGI "Cordero de Navarra"[16] and for its conventional counterpart. In what follows, we will use the same notation as in the previous case study for the prices to be analysed so we have: FPI and RPI for farm and retail prices of the PGI lamb respectively, and FP and RP for farm and retail prices of the conventional lamb, respectively.

Nominal prices are illustrated in Figure 3, in which the conventional lamb is shown in Panel A and the PGI lamb in Panel B. Prices of conventional lamb seem to suggest a co-movement over time with more fluctuation periods at retail level and lagged responses in some periods of farm prices after changes in retail prices. In Panel B, again, the two price series appear to obey a relationship over the sample. As in the conventional system, we can easily identify more volatile episodes at retail level than at farm level. From visual inspection of the two panels, prices display sufficient volatility and an increasing pattern over time. All the price series were transformed into natural logs according to theory.

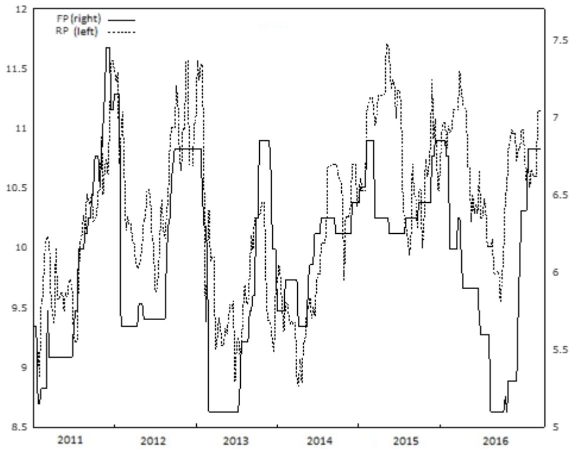
A summary of basic statistics of all of the price series are reported in Table 8. The statistics indicate that the trend, though extremely small, is only significant in the conventional system not in the quality system. All the prices seem to strongly reject the null hypothesis of no ARCH effects.

Unit root tests and stationarity tests are applied to analyse the order of integration of all the series. Results reported in Table 9 point out that the series can be characterized as non-stationary, that is, integrated of order 1. The presence of a unit root in all the series is robust to the presence of possible structural changes according to Carrion-i-Silvestre et al (2009).

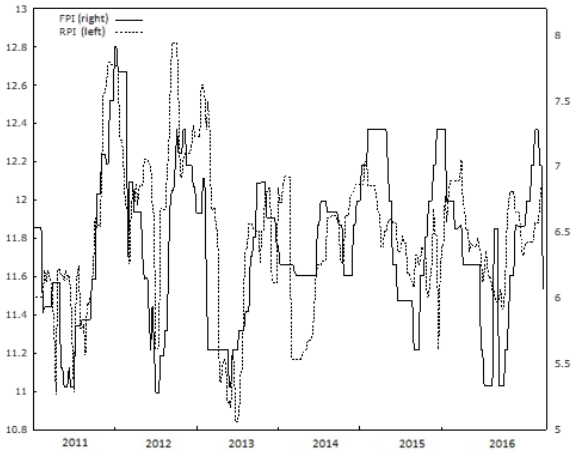
[15] "Nafarroako Arkumea" in Euskera.
[16] In the remainder of the section we use indistinctly light lamb and lamb.

Figure 3. Evolution of Spanish lamb sector price series

Panel A. Conventional "Cordero de Navarra"



Panel B. PGI "Cordero de Navarra"



Source: Own calculation based on Regional Government of Navarra (Spain), Observatory of Agricultural Prices database. Vertical axes are measured in €/kg carcass.

Since all the price series are nonstationary, we can assess whether there is co-integration between each pair of prices in each system. As our co-integration testing approach is based on the unrestricted VAR model, we first estimate the lag order which ensures the presence of no autocorrelation in the system. Results are reported in Table 10.

Based on the lag order choice for the two VAR model specifications, we test for co-integration rank following Johansen (2002)[17]. Table 11 shows the results and the respective long-run relationships for each system. The coefficients can also be interpreted as price elasticities since the prices have been transformed into logs. In the conventional system, there is a strong positive relationship (42.3%) which implies that an increase in farm prices will lead to an increase in

[17] The use of the approach of Engle and Granger (1987) also concludes in favour of the presence of co-integration relationships in each system.

retail prices. For the quality system, the results imply a slightly weaker relationship between prices, with around 20% of price elasticity.

Table 8. Summary of descriptive statistics for the Spanish lamb price series

	FP	RP	FPI	RPI
Mean	6.080	10.301	6.380	11.818
Median	6.150	10.299	6.321	11.842
Minimum	5.100	8.843	5.285	10.840
Maximum	7.450	11.703	7.912	12.822
Standard deviation	0.580	0.685	0.606	0.393
Skewness	-0.058	-0.033	0.110	0.050
Kurtosis (excess)	-0.899***	-0.905***	-0.620**	0.240
Jarque-Bera test	10.576***	10.602***	5.622*	0.881
Engle (1982)'s test	257.347***	187.593***	229.903***	221.897***
Trend	$7.059e^{-04}$*	$2.489e^{-03}$***	$5.059e^{-03}$	$-1.907e^{-04}$
# observations	312	312	312	312

Notes: We have considered logarithmic transformations of the prices in our application. The use of the test of Engle (1982) allows us to check whether there are ARCH effects. In this case, we have used 2 lags. ***, ** and * denote statistically significant at 1%, 5% and 10% level of significance.

Table 9. Results of the integration order for the Spanish lamb price series

Panel A. Unit root and stationarity tests

	MSB	KPSS
FP	0.185 (1)	0.680 (0)**
RP	0.184 (5)	1.189 (0)**
FPI	0.155 (1)	0.319 (1)**
RPI	0.158 (1)	0.349 (1)**

Panel B. Unit root tests allowing for structural breaks (Carrion-i-Silvestre et al, 2009)

	ADF	MSB	ℓ	Tb
FP	-2.521 (-3.347)	0.195 (0.146)	0	2011:47 (47)
RP	-3.246 (-3.451)	0.147 (0.142)	5	2012:52 (104)
FPI	-2.950 (-3.378)	0.173 (0.145)	0	2012:05 (57)
RPI	-4.022 (-4.142)	0.131 (0.119)	0	2013:03 (107) 2013:41 (145)

Notes: In Panel A, we apply the MSB unit root test as in Ng and Perron (2001) and the KPSS stationarity test. The truncation lag parameter, k, presented in parentheses and is estimated using the MAIC. ** denotes statistically significant at 5% level of significance since the asymptotic critical values at 5% level for the constant (trend) case are respectively 0.233 (0.168) for the MSB, and 0.463 (0.146) for the KPSS. As for the prices in the conventional system we have applied the tests in their respective trend version. In Panel B, ℓ reflects the bandwidth parameter for the KPSS test selected with the automatic bandwidth procedure of Andrews (1991) for the kernel-based estimator of the long-run variance. The critical value at the 5%

significance level of each test is shown in parentheses. Tb reflects the time breaks, that is, the date when a structural break was endogenously detected and the corresponding number of observation is presented in parentheses.

Table 10. VAR lag-length selection for the Spanish lamb sector

Panel A. Conventional System

IC	k
BIC	1
AIC	6
HQ	1

Panel B. Quality System

IC	k
BIC	1
AIC	1
HQ	1

Notes: k denotes the number of lags of the unrestricted VAR model. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Table 11. Results of the Co-integration analysis for the Spanish lamb sector

Conventional System: RP – FP

Rank	Eigen value	λ^{*}_{trace}	p-value
0	0.073	24.494	0.001
1	0.021	6.328	0.172

Co-integration relationship: $ECT = RP_t - 1.568^{***} - 0.423^{***}FP_t$

Quality System: RPI - FPI

Rank	Eigen value	λ^{*}_{trace}	p-value
0	0.079	34.032	0.000
1	0.027	8.450	0.069

Co-integration relationship: $ECTI = RPI_t - 2.103^{***} - 0.198^{***}FPI_t$

Notes: *** denotes statistically significant at 1% level of significance. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Table 12 reports the estimated results obtained from the estimation of the conditional mean model. First, we can observe that the retail price in the conventional system reverts slower (by 6%) to the long-run equilibrium after deviations than the retail price in the quality system, by 13%. Again, the adjustments in the short-run are faster in the quality system than in the conventional system as one could expect. Conventional retail prices are only affected by the farm price first lag. Conventional farm prices are affected by its own second lag and the retail second lag. As to the quality system, we cannot find statistical significance of any lags. Correct model specification is checked by carrying out a multivariate residual analysis. The lower panel of Table 12 reports the results for the multivariate Q statistic and ARCH LM tests for each system. According to this, we can see that volatility is time-varying and no autocorrelation exists.

Table 12. Results for the conditional mean model for the Spanish lamb sector

Short-run parameters for the Conventional System

$$\begin{pmatrix}\Delta RP_t\\ \Delta FP_t\end{pmatrix} = \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix} ECT_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121}\\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-1}\\ \Delta FP_{t-1}\end{pmatrix} + \begin{pmatrix}\delta_{112} & \delta_{122}\\ \delta_{212} & \delta_{222}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-2}\\ \Delta FP_{t-2}\end{pmatrix} + \varepsilon_t$$

Short-run parameters for the Quality System

$$\begin{pmatrix}\Delta \mathrm{RPI_t}\\ \Delta \mathrm{FPI_t}\end{pmatrix} = \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix} \mathrm{ECTI}_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121}\\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta \mathrm{RPI}_{t-1}\\ \Delta \mathrm{FPI}_{t-1}\end{pmatrix} + \varepsilon_t$$

	Conventional System		*Quality System*	
	$i = 1$	$i = 2$	$i = 1$	$i = 2$
α_i	-0.062** (0.018)	-0.006 (0.021)	-0.132*** (0.057)	0.099 (0.075)
δ_{11i}	-0.035 (0.060)	-0.052 (0.042)	0.076 (0.057)	-
δ_{12i}	0.143*** (0.048)	0.024 0.049)	0.015 (0.023)	-
δ_{21i}	-0.105 (0.077)	0.037 (0.051)-	0.059 (0.149)	-
δ_{22i}	0.144** (0.061)	0.076 (0.051)	0.016 (0.060)	-
Hosking's test: 39.15 (0.82)			Hosking's test: 43.391 (0.66)	
Multivariate ARCH LM test: 30.57*** (0.00)			Multivariate ARCH LM test: 53.62*** (0.00)	

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for multivariate tests in which contains the p-value. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

We then estimate the conditional variance model using the asymmetric multivariate BEKK-GARCH approach. We show the results in Table 13 together with usual tests on standardized residuals to guarantee modelling adequacy. We checked that the eigenvalues were less than unity, which implies that the two estimated models are covariance stationary. By applying the Nyblom (1989) test, we ensure the stability of the models, at least at the 5% of significance level, for the conventional and quality systems, respectively. In both systems, we test whether the presence of volatility is time-varying. Thus, the null hypothesis that parameters in matrices A, B, and D are equal to zero is strongly rejected. Asymmetric effects are found in the two systems, though slightly weaker in the quality system. The null that the parameters in matrix D are equal to zero is rejected.

Figure 4 illustrates the predicted conditional variances for the conventional (left panel) and quality system (right panel). We can see that volatilities are not constant over time and are slightly higher in the conventional system, as seen when comparing the magnitudes of the respective predicted volatilities in the left- and right-hand side panels of Figure 4. Volatility at the farm level in the conventional system is characterized by rises in the first weeks of 2012 and 2013. This may be interpreted as volatility spill-overs from the retail level. At the farm level, sharp rises are observed in the first weeks of year 2012, but they are much less volatile in 2013. In the quality system, the volatility in retail prices is characterized by appreciable fluctuations exhibiting a downward trend by year 2013 which ends in 2016 with an abrupt rise before starting to decrease. Magnitudes are slightly higher than at the farm price level, in which we can observe a similar pattern. We can infer volatility spill-overs from retail to the farm level. Finally, conditional correlations suggest a quite strong relationship between market shocks in the conventional system with appreciable fluctuations from positive to negative, which can be interpreted as a result of the limited volatility spilled over from retail prices. However, the conditional correlations in the quality system are small but persistent.

Table 13. Results for the conditional variance model for the Spanish lamb sector

GARCH model parameters

Asymmetric BEKK for the Conventional System

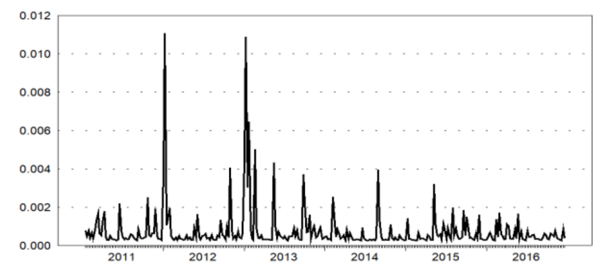
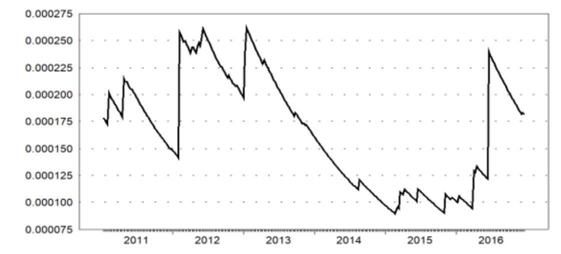
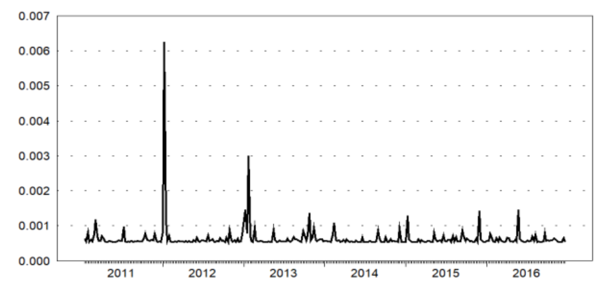
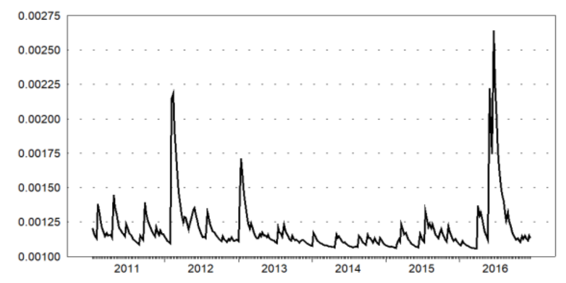
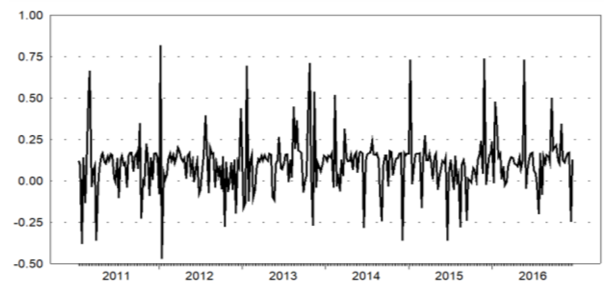
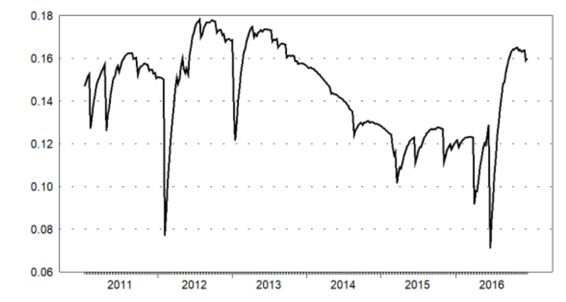
$$C = \begin{pmatrix} c_{11} & 0 \\ c_{21} & c_{22} \end{pmatrix}, A = \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix}, B = \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \text{ and } D = \begin{pmatrix} d_{11} & d_{12} \\ d_{21} & d_{22} \end{pmatrix}$$

	Conventional System		Quality System	
	$i = 1$	$i = 2$	$i = 1$	$i = 2$
c_{1i}	0.014 *** (0.000)	-	-0.000 (0.001)	-
c_{2i}	0.004 (0.003)	0.023 *** (0.001)	-0.016*** (0.002)	0.000 (0.000)
a_{1i}	0.297*** (0.082)	0.148* (0.085)	-0.004 (0.014)	-0.125 (0.111)
a_{2i}	-0.483*** (0.041)	0.224*** (0.051)	0.003 (0.012)	0.167* (0.091)
b_{1i}	-0.347*** (0.092)	0.081 (0.059)	0.994*** (0.006)	0.046 (0.037)
b_{2i}	0.232*** (0.066)	0.117 (0.343)	0.000 (0.011)	0.874*** (0.007)
d_{1i}	0.257* (0.144)	0.279** (0.103)	-0.002 (0.014)	0.035 (0.055)
d_{2i}	0.690*** (0.051)	0.084 (0.081)	0.055*** (0.022)	-0.009 (0.059)

LR test for the null of joint significance of parameters of matrices A,B,D: 32.069*** (0.00)		298704.23*** (0.00)
LR test for the null of joint significance of parameters of matrices A,B: 30.176*** (0.00)		88475.78***(0.00)
LR test for the null of asymmetric effects:	57.400*** (0.00)	3.67*** (0.00)
LR test for BEKK cross effects:	40.411*** (0.00)	2.185** (0.04)
Nyblom (1989) joint stability test:	4.515 (0.46)	3.413* (0.07)
Stable roots of the system?	YES	YES

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for the multivariate tests reported in the lower panel in which contains the p-value. We have computed the simplexalgorithm to obtain the initial conditions together with the BFGS (Broyden, Fletcher, Goldfarb, Shanno) algorithm is then employed to obtain the final estimate of the variance-covariance matrix and the corresponding standard errors. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Figure 4. Predicted volatilities and conditional correlations for the Spanish lamb sector

Conventional System	*Quality System*
Predicted retail volatility	
	
Predicted farm volatility	
	
Conditional correlation	
	

Source: Own source from results obtained with CATS (Dennis et al, 2006) in RATS 9.0.

As the individual coefficients derived from the estimation of the conditional variance model cannot be directly interpreted, we can draw conclusions from the conditional variance equations presented in Table 14.

For the conventional system, we can see that the volatility in the retail price is not only directly affected by its own past volatility (h_{11t-1}), past farm volatility (h_{22t-1}) but also indirectly through the covariance term (h_{12t-1}). This indirect effect implies that the high correlation between retail and farm level is quite important to reduce price instability, as well as safeguarding the stability when prices move in the same direction. This observation may also be interpreted as evidence of retail power having an effect on instability in producer markets. Past shocks to farm markets (u^2_{2t-1}) have an asymmetric effect on retail price and farm price (h_{22t}) volatilities which may imply more sensitivity to price decreases than increases. However, the past correlation between retail and farm has only significant effects on retail price volatility (h_{11t}).

For the quality system, the results suggest that retail price volatility (h_{11t}) is only affected by its own past volatility (h_{11t-1}), and it seems that the correlation between retail and farm levels has no impact on price volatility (h_{12t-1} is not statistically significant). Farm price volatility (h_{22t}) is only directly influenced by its own lagged volatility because past shocks seem not statistically significant.

Table 14. Conditional variance equations for the Spanish lamb sector

Conventional System

$$h_{11} = 2.29\,e^{-04^{***}} + 0.122^{*}h_{11t-1} + 0.047^{*}\mathrm{h}_{22t-1} - 0.151^{**}h_{12t-1} + 0.088\,u_{1t-1}^{2} + 0.232^{***}\,u_{2t-1}^{2} - 0.285^{**}u_{1t-1}u_{2t-1} + 0.062\,v_{1t-1}^{2} + 0.474^{***}v_{2t-1}^{2} + 0.344^{***}v_{1t-1}v_{2t-1}$$

$$h_{22} = 5.14e^{-04^{***}} + 6.95\,e^{-03}h_{11t-1} + 0.017\,\mathrm{h}_{22t-1} + 0.022\,h_{12t-1} + 0.022\,u_{1t-1}^{2} + 0.050^{*}u_{2t-1}^{2} + 0.066\,u_{1t-1}u_{2t-1} + 0.077\,v_{1t-1}^{2} + 0.007\,v_{2t-1}^{2} + 0.047\,v_{1t-1}v_{2t-1}$$

Quality system

$$h_{11} = 2.44e^{-04} + 0.988^{***}h_{11t-1} + 0.000\,\mathrm{h}_{22t-1} + 3.07\,e^{-04}\,h_{12t-1} + 1.4\,e^{-05}\,u_{1t-1}^{2} + 7.0\,e^{-06}\,u_{2t-1}^{2} - 1.9\,e^{-05}\,u_{1t-1}u_{2t-1} + 6.0\,e^{-06}\,v_{1t-1}^{2} + 3.03\,e^{-03}v_{2t-1}^{2} - 2.61\,e^{-04^{**}}v_{1t-1}v_{2t-1}$$

$$h_{22} = 0.000 + 0.002\,h_{11t-1} + 0.765^{***}\,\mathrm{h}_{22t-1} + 0.081\,h_{12t-1} + 0.016\,u_{1t-1}^{2} + 0.028\,u_{2t-1}^{2} - 1.9\,e^{-05}\,u_{1t-1}u_{2t-1} + 1.24\,e^{-03}\,v_{1t-1}^{2} + 7.8\,e^{-05}v_{2t-1}^{2} - 6.21\,e^{-04}v_{1t-1}v_{2t-1}$$

Notes: h_{11} retail price, h_{22} farm price variance. Estimated parameters of h_{ijt-1},$i,j = 1,2$ reflects direct and indirect volatility transmission between prices, whereas those of u_{it-1}^{2} and $u_{it-1}u_{jt-1}$ indicates how price volatility is affected by markets shocks. ***, ** and * denotes statistical significance at the 1%, 5% and 10% level, respectively. "e" indicates scientific notation (with exponent).

To sum up, our results provide evidence on significant volatility spill-overs along the two lamb marketing chains. Also, we have found that retailers exert to some extent market power which may result in more stability in prices. As in the previous case study, the patterns are found to be in magnitude smaller when compared to those in the conventional system.

3.3. The Italian cheese sector

We follow a similar structure as in the two preceding cases. After a sound overview of the quality designated product, we present and discuss the results for the pair-wise analysis in which we compare the price volatility results for the quality product and its conventional counterpart.

3.3.1. An overview of the PDO "Parmigiano Reggiano"

Parmigiano Reggiano is one of the most valued cheeses with a long tradition and history, which goes back to ancient times, circa 1200 in the Benedictine monasteries settled close to the river Po and the Apennines where it was first produced. Undoubtedly, it can be said that this product is strictly linked not only to the territory but also to all the people who manufacture it, on whose knowledge the product crucially depends.

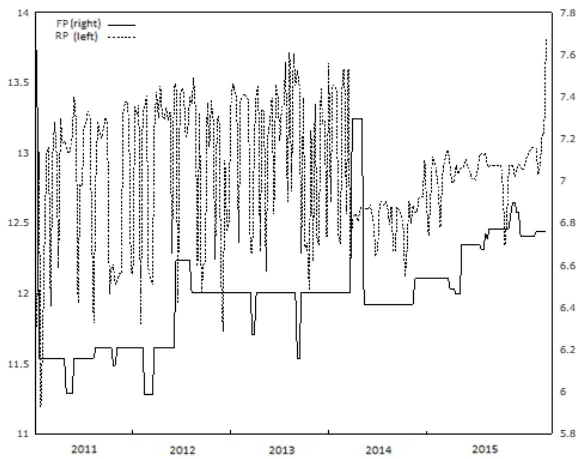
Parmigiano Reggiano is designated as a Protected Designation of Origin (PDO) product to safeguard producers and consumers, and guarantee the distinctive features of the cheese as well as its linkages to the zone of origin. Its production is carried out according to the PDO specifications reflected in the respective bid specifications and strictly controlled by official institutions. Furthermore, the appellation PDO Parmigiano Reggiano can only designate cheeses produced and processed in the place of origin, and manufactured according to strict standards that require precise production methods, controlled feeding of cows, and qualitative selection and designation (Parmigiano Reggiano official website, 2012).

3.3.2. Data and Results

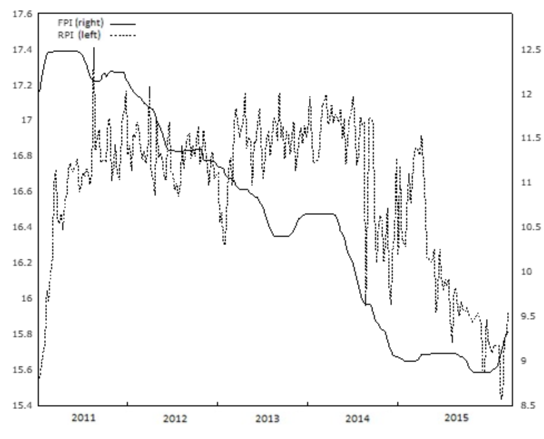
Parmigiano Reggiano (18-24 months) prices, expressed in Euros, at the farm (FPI) and retail (RPI) levels are observed weekly from 2011 to 2015. Prices for the Generic Caciotta cheese (the conventional counterpart considered for this study) are also observed at farm (FP) and retail (RP) levels for the same period. All of the four series are obtained from the Istituto di Servizi per il Mercato Agricolo Alimentare (ISMEA) database. Nominal prices are illustrated in Figure 9, in which Panel A illustrates the Generic Caciotta cheese and Parmigiano Reggiano in Panel B. In both panels, prices seem to obey a long-term relationship, with the retail prices being much more volatile than farm prices. Furthermore, all the prices exhibit visible fluctuations along with a positive growth in the case of the conventional cheese and a decreasing pattern for the quality cheese.

Figure 5. Evolution of the Italian cheese sector price series

Panel A. Generic Caciotta



Panel B. PDO Parmigiano Reggiano



Source: Own calculation based on ISMEA database. Vertical axes are measured in €/kg.

Table 15 reports a summary of the descriptive statistics of the series for each system. The trend is significant for RP, and FPI and RPI but negative and rather small. Non-normality and ARCH effects are found in all the series. As before, we examined whether the logged price series contain a unit root and it can be concluded that all the series are integrated of order 1 (Table 16).

Table 15. Summary of descriptive statistics for Italian cheese price series

	FP	RP	FPI	RPI
Mean	6.45	12.860	10.654	16.605
Median	6.473	12.910	10.655	16.752
Minimum	5.988	11.192	8.875	15.435
Maximum	7.620	13.824	12.485	17.412
Standard deviation	0.253	0.473	1.222	0.419
Skewness	0.988***	-0.503***	-0.068	-0.948***
Kurtosis	2.733***	-0.298	-1.328***	-0.138
Jarque-Bera test	123.240***	11.945***	19.300***	39.139***
Engle (1982)'s test	220.225***	20.924***	257.906***	208.045***
Trend	-7.682e^{-05}	0.002***	-0.003***	-0.016***
# observations	260	260	260	260

Notes: We have considered logarithmic transformations of the prices in our application. The use of the test of Engle (1982) allows us to check whether there are ARCH effects. In this case, we have used 2 lags. *** denotes statistically significant at 1% level.

According to theory, co-integration exists when nonstationary variables, in our case prices, show a tendency to move together in the long-run and deviations from this equilibrium due to unexpected shocks tend to revert eventually. Also, co-integration analysis is based on the unrestricted vector autoregressive (VAR) model. See Juselius (2006) for an excellent illustration of the co-integrated VAR model.

Prior to testing for co-integration we select the number of lags of the unrestricted VAR model. Table 3 reports the results.

Once the lag order has been determined for each system, we test for co-integration and select the co-integration rank using the Bartlett corrected trace test λ^{*}_{trace} as in Johansen (2002)[18]. Results are presented in Table 4. Recall that prices are considered in logs. For this reason, we can interpret the co-integration parameters as price elasticities. Subsequently, we find a positive relationship for the protected system (11%) implying that an increase in farm prices will lead to a rise in retail prices. However, for the conventional system, the relationship seems only significant at 20% with low influence of farm prices on retail prices, around 2%.

[18] Note that if we apply the method suggested in Engle and Granger (1987) we also found the existence of respective co-integration relationships in the two systems.

Table 16. Results of the integration order for the Italian cheese price series

Panel A. Unit root and stationarity tests

	MSB	KPSS
FP	0.248 (1)	0.404 (0)**
RP	0.353 (10)	0.759 (0)**
FPI	0.247 (1)	2.656 (0)**
RPI	0.253 (3)	2.820 (0)**

Panel B. Unit root tests allowing for structural breaks (Carrion-i-Silvestre et al, 2009)

	ADF	MSB	ℓ	Tb
FP	-2.775 (-3.312)	0.300 (0.161)	1	2014:27 (183)
RP	-2.041 (-3.289)	0.201 (0.160)	10	2011:27 (27)
FPI	-2.342 (-3.835)	0.224 (0.129)	7	2013:5 (107); 2014:3 (169)
RPI	-2.216 (-3.443)	0.229 (0.143)	3	2013:4 (106)

Notes: In Panel A, we apply the MSB unit root test as in Ng and Perron (2001) and the KPSS stationarity test. The truncation lag parameter, k, presented in parentheses and is estimated using the MAIC. Also, ** denotes statistically significant at 5% level of significance since the asymptotic critical values at 5% level for the constant (trend) case are respectively 0.233 (0.168) for the MSB, and 0.463 (0.146) for the KPSS. In Panel B, ℓ reflects the bandwidth parameter for the KPSS test selected with the automatic bandwidth procedure of Andrews (1991) for the kernel-based estimator of the long-run variance. The critical value at the 5% significance level of each test is shown in parentheses. Tb reflects the time breaks, that is, the date when a structural break was endogenously detected and the corresponding number of observation is presented in parentheses.

Table 17. VAR lag-length selection for the Italian cheese sector

Panel A. Conventional System

IC	**lags**
BIC	1
AIC	1
HQ	1

Panel B. Quality System

IC	**lags**
BIC	2
AIC	5
HQ	3

Notes: Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Table 18. Results of the Co-integration analysis for the Italian cheese sector

Conventional System: RP – FP

Rank	Eigen value	λ^{*}_{trace}	p-value
0	0.208	68.082	0.000
1	0.035	8.995	0.053
Co-integration relationship: $ECT = RP_t - 2.586^{***} + 0.018\, FP_t$			
Quality System: RPI - FPI			
Rank	Eigen value	λ^{*}_{trace}	p-value
0	0.082	28.699	0.002
1	0.029	7.207	0.119
Co-integration relationship: $ECTI = RPI_t - 2.554^{***} - 0.108^{***}\, FPI_t$			

Notes: *** denotes statistically significant at 1% level of significance. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

The asymmetric multivariate BEKK-GARCH model is jointly estimated utilizing maximum likelihood methods. The results obtained for the conditional mean model are shown in Table 19 whilst those for the conditional volatility model are presented in Table 20.

Results in Table 19 suggest for the conventional system that retail prices and farm prices respond to deviations from the long-run equilibrium by reverting to it at 57% and 2% respectively. For the quality system, the response is taken by farm prices at 1%. Faster adjustments are found in the conventional system. Moreover, conventional retail prices are affected by own second and third lagged values and farm lags. In the quality system, retail prices are only affected by own first and second lagged values. Farm conventional prices exhibit no significant dependence of any lagged variables, but in the quality system, farm premium prices only depend on its own lags. Multivariate residual analysis suggests model adequacy, as can be seen from the outcome of the multivariate Q statistic of Hosking (1980, 1981) and the multivariate ARCH LM tests for each system reported in the lower panel of Table 5. The latter strongly rejects the null of no ARCH effects implying time-varying volatility in both systems.

Asymmetric BEKK-MGARCH parameter estimates are reported in Table 20. Correct model specification is ensured by the stability of the roots for each system and by using the Nyblom (1989) test. The presence of time-varying volatility is guaranteed with the rejection of the null hypothesis that parameters in matrices A, B, and D are equal to zero. Asymmetric effects are found in both systems as the null hypothesis that parameters in matrix D are equal to zero is strongly rejected.

Table 19. Results for the conditional mean model for the Italian cheese sector

Short-run parameters for the Conventional System

$$\begin{pmatrix}\Delta RP_t\\ \Delta FP_t\end{pmatrix} = \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix} ECT_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121}\\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-1}\\ \Delta FP_{t-1}\end{pmatrix} + \begin{pmatrix}\delta_{112} & \delta_{122}\\ \delta_{212} & \delta_{222}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-2}\\ \Delta FP_{t-2}\end{pmatrix} + \begin{pmatrix}\delta_{113} & \delta_{123}\\ \delta_{213} & \delta_{223}\end{pmatrix}\begin{pmatrix}\Delta RP_{t-3}\\ \Delta FP_{t-3}\end{pmatrix} + \varepsilon_t$$

Short-run parameters for the Quality System

$$\begin{pmatrix}\Delta \mathrm{RPI}_t\\ \Delta \mathrm{FPI}_t\end{pmatrix} = \begin{pmatrix}\alpha_1\\ \alpha_2\end{pmatrix} \mathrm{ECTI}_{t-1} + \begin{pmatrix}\delta_{111} & \delta_{121}\\ \delta_{211} & \delta_{221}\end{pmatrix}\begin{pmatrix}\Delta \mathrm{RPI}_{t-1}\\ \Delta \mathrm{FPI}_{t-1}\end{pmatrix} + \begin{pmatrix}\delta_{112} & \delta_{122}\\ \delta_{212} & \delta_{222}\end{pmatrix}\begin{pmatrix}\Delta \mathrm{RPI}_{t-2}\\ \Delta \mathrm{FPI}_{t-2}\end{pmatrix} + \varepsilon_t$$

	Conventional System			*Quality System*	
	$i=1$	$i=2$	$i=3$	$i=1$	$i=2$
α_i	-0.567*** (0.024)	-0.020*** (0.003)	-	-0.003 (0.016)	-0.009** (0.004)
δ_{11i}	-0.044 (0.067)	-0.199*** (0.046)	-0.079* (0.041)	-0.244* (0.058)	-0.187* (0.053)
δ_{12i}	0.035 (0.029)	0.100*** (0.022)	0.104*** (0.018)	0.280 (0.353)	0.186 (0.464)
δ_{21i}	0.006 (0.042)	0.001 (0.001)	-0.002 (0.020)	0.006 (0.011)	0.010 (0.011)
δ_{22i}	-0.072 (0.071)	-0.135 (0.124)	-0.019 (0.031)	0.611* (0.085)	0.200* (0.098)
Hosking's test: 53.574 (0.27)				Hosking's test: 37.306 (0.87)	
Multivariate ARCH LM test: 28.67*** (0.00)				Multivariate ARCH LM test: 30.97*** (0.00)	

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for multivariate tests in which contains the p-value. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Figure 6 illustrates the respective predicted conditional variances for the conventional system in the left-hand panel and for the quality system in the right-hand panels. From the conventional graphs we can see that fluctuations are higher in retail prices than in farm prices, which may be read as volatility spill-overs from the retail level. Also, this can be seen if compared the magnitudes of the respective predicted volatilities in the left- and right-hand side panels of Figure 6. Fluctuations, mainly at retail level, tend to reduce significantly by 2014 and 2015. For the quality system, predicted conditional variances are lower than those illustrated in the left-hand panel, especially for the farm level. Moreover, conditional correlations are slightly superior in the conventional system than in the quality system. In both cases, we can observe significant changes from positive to negative (more often in the quality system) implying a low transmission of price signals or unanticipated responses to negative values, which can be also interpreted as reduced volatility spilled over from retail markets.

We draw conclusions from the conditional variance equations as it is not possible to directly interpret the estimated coefficients from the BEKK-GARCH model. Results reported in Table 21 suggest that conventional retail price volatility is influenced by its own past volatility (h_{11t-1}), and its own shocks (u^2_{1t-1}). The covariance term (h_{12t-1}) is not relevant, and this may imply that the decrease in price instability cannot be guaranteed and hence, that the retail market does not exert much influence on farm prices. Asymmetric effects are only found to be significant for farm price volatility together with the correlation of past shocks between retail and farm prices which may reduce farm price volatility.

Table 20. Results for the conditional variance model for the Italian cheese sector

GARCH model parameters

Asymmetric BEKK for the Conventional System

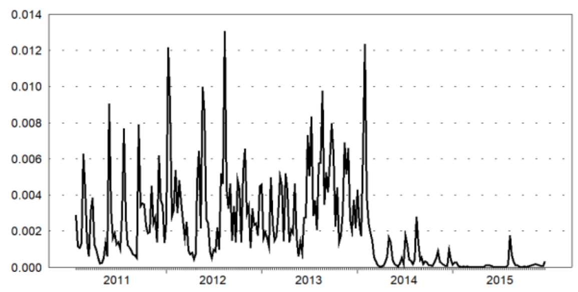
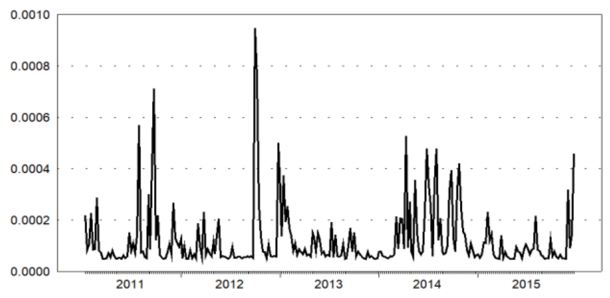
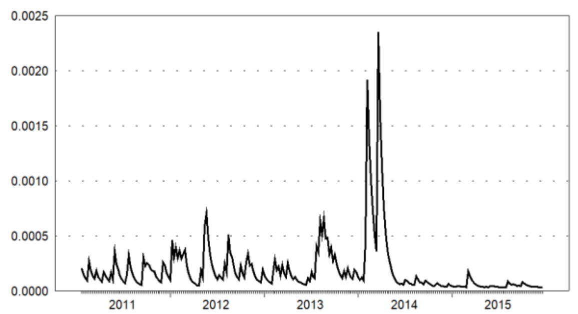
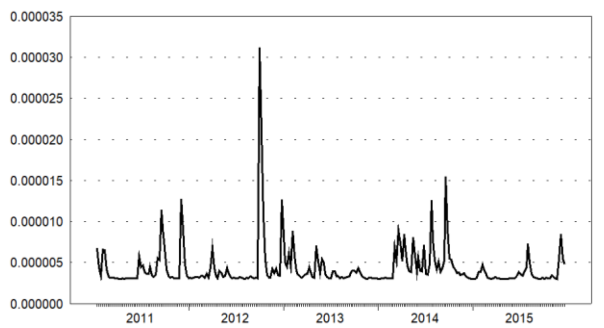
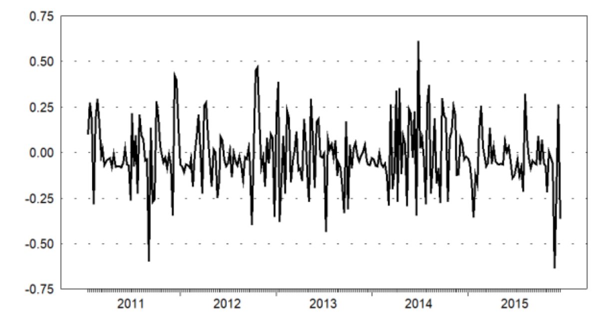
$$C = \begin{pmatrix} c_{11} & 0 \\ c_{21} & c_{22} \end{pmatrix}, A = \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix}, B = \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \text{ and } D = \begin{pmatrix} d_{11} & d_{12} \\ d_{21} & d_{22} \end{pmatrix}$$

	Conventional System		Quality System	
	$i = 1$	$i = 2$	$i = 1$	$i = 2$
c_{1i}	-0.002 (0.001)	-	0.005*** (0.001)	-
c_{2i}	0.000 (0.002)	0.003*** (0.001)	-0.001*** (0.000)	-4.08 e-^{07} (0.000)
a_{1i}	1.078*** (0.129)	0.011 (0.039)	-0.423*** (0.060)	-0.019 (0.012)
a_{2i}	-0.116 (0.131)	0.342** (0.152)	-3.076*** (0.727)	0.281* (0.159)
b_{1i}	0.553*** (0.061)	-0.039 (0.036)	0.095 (0.201)	-0.032 (0.034)
b_{2i}	0.063 (0.059)	0.820*** (0.039)	2.975*** (0.460)	0.606*** (0.109)
d_{1i}	-0.317* (0.175)	-0.204*** (0.046)	-0.036 (0.036)	-0.013 (0.009)
d_{2i}	0.016 (0.041)	0.097 (0.161)	0.493 (0.584)	0.415*** (0.132)
LR test for the null of joint significance of parameters of matrices A,B,D: 433.924*** (0.00)			52.055*** (0.00)	
LR test for the null of joint significance of parameters of matrices A,B: 262.707*** (0.00)			36.545*** (0.00)	
LR test for the null of asymmetric effects:		11.600*** (0.00)	3.991*** (0.00)	
LR test for BEKK cross effects:		9.360*** (0.00)	19.601*** (0.00)	
Nyblom (1989) joint stability test:		7.106** (0.02)	5.126 (0.20)	
Stable roots of the system?		YES	YES	

Source: ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively. Standard errors in parenthesis, except for the multivariate tests reported in the lower panel in which contains the p-value. We have computed the simplex algorithm to obtain the initial conditions together with the BFGS (Broyden, Fletcher, Goldfarb, Shanno) algorithm is then employed to obtain the final estimate of the variance-covariance matrix and the corresponding standard errors. Results are obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Premium retail price volatility (h_{11t}) is affected by the past farm price volatility (h_{22t-1}), together with the past shocks originated at retail and farm levels (both coefficients are statistically significant). Again, the correlation between retail and farm levels has no impact on price volatility (h_{12t-1} is not statistically significant). Farm price volatility is directly affected by its own past volatility (h_{22t-1}) and by the strong correlation of past shocks at market. On the other hand, asymmetric effects are not significant. It seems that past shocks originated at retail markets do affect retail and farm volatilities (u^2_{1t-1} is statistically significant) by increasing market volatility.

Figure 6. Predicted volatilities and conditional correlations for the Italian cheese sector

<table>
<tr><th>Conventional System</th><th>Quality System</th></tr>
<tr><td colspan="2">Predicted retail volatility</td></tr>
<tr><td>
</td><td>
</td></tr>
<tr><td colspan="2">Predicted farm volatility</td></tr>
<tr><td>
</td><td>
</td></tr>
<tr><td colspan="2">Conditional correlation</td></tr>
<tr><td>
</td><td>
</td></tr>
</table>

Source: Own source from results obtained with CATS (Dennis et al, 2006) in RATS 9.0.

Similar to the two previous case studies, price patterns along these two supply chains are found to be characterized with volatility spill-overs, which are more significant in the conventional chain. Moreover, retailers are found not to exert much market power, and past shocks in the premium system are found to have no significant asymmetric effect on price volatility in the two markets (coefficients are not statistically significant) which may lead to responses equally sensitive to price decreases and to price increases.

Table 21. Conditional variance equations for the Italian cheese sector

Conventional System

$$h_{11} = 4.0e^{-06} + 0.306^{***}h_{11t-1} + 0.004\ \mathrm{h}_{22t-1} + 0.070\ h_{12t-1} + 1.161^{***}\ u_{1t-1}^2 + 0.014\ u_{2t-1}^2 - 0.25\ u_{1t-1}u_{2t-1} + 0.100\ v_{1t-1}^2 + 2.46\ e^{-04}v_{2t-1}^2 - 0.010\ v_{1t-1}v_{2t-1}$$

$$h_{22} = 9.0\ e^{-06} + 1.5\ e^{-03}h_{11t-1} + 0.672^{***}\ \mathrm{h}_{22t-1} - 0.064\ h_{12t-1} + 1.17\ e^{-04}\ u_{1t-1}^2 + 0.117\ u_{2t-1}^2 - 0.251^{***}\ u_{1t-1}u_{2t-1} + 0.042^{***}v_{1t-1}^2 + 0.010\ v_{2t-1}^2 - 0.040\ v_{1t-1}v_{2t-1}$$

Quality system

$$h_{11} = 2.2e^{-05^{***}} + 0.009\ h_{11t-1} + 8.853^{***}\ \mathrm{h}_{22t-1} + 0.565\ h_{12t-1} + 0.179^{***}\ u_{1t-1}^2 + 9.464^{**}\ u_{2t-1}^2 + 2.602^{***}\ u_{1t-1}u_{2t-1} + 0.001\ v_{1t-1}^2 + 0.243\ v_{2t-1}^2 - 0.036\ v_{1t-1}v_{2t-1}$$

$$h_{22} = 0.000 + 0.001\ h_{11t-1} + 0.367^{***}\ \mathrm{h}_{22t-1} - 0.038\ h_{12t-1} + 3.52e^{-04}\ u_{1t-1}^2 + 0.0787\ u_{2t-1}^2 + 2.602^{***}\ u_{1t-1}u_{2t-1} + 1.62e^{-04}\ v_{1t-1}^2 + 0.173\ v_{2t-1}^2 - 0.011\ v_{1t-1}v_{2t-1}$$

Notes: h_{11} retail price, h_{22} farm price variance. Estimated parameters of h_{ijt-1}, $i,j = 1,2$ reflects direct and indirect volatility transmission between prices, whereas those of u_{it-1}^2 and $u_{it-1}u_{jt-1}$ indicates how price volatility is affected by markets shocks. ***, ** and * denotes statistical significance at the 1%, 5% and 10% level, respectively. "e" indicates scientific notation (with exponent).

4. CONCLUDING REMARKS

This paper provides insights in the analysis of price volatility along the marketing chain of food products protected with European food quality schemes. We focused on food quality products because of the growing interest of consumers, producers and food policy makers in the potential of these schemes to increase competitiveness in both domestic and international markets and maintain social and territorial cohesion.

To date, literature investigating price relationships has mainly focused on price transmission for conventional products but, in the last few years, studies have started to investigate how price volatility is transmitted along the food supply chain. However, there is a gap in the literature since food quality schemes have not been examined yet. Our study contributes to fill this gap and provides novel results on price volatility relationships along the supply chain for FQS designated goods.

Three case-studies integrate this analysis. The first and second studies deal with the Spanish meat sector whereas the third study focuses on the Italian cheese sector. We examine the PGI beef from Navarra, the PGI lamb from Navarra, and the PDO Parmigiano Reggiano. To provide a better understanding, the results derived from these systems are compared to their conventional counterparts to assess the differences in the transmission of price volatilities along the supply chain.

We work with weekly prices observed at two stages of the supply chain for each system, farm and retail levels, covering a period of several years after the rise in prices that took place in 2007/2008, and the second rise in the beginning of 2011. To achieve our main objective, we have estimated VECM and multivariate BEKK-GARCH models allowing for asymmetric effects in response. A summary of the results is presented below.

Results derived from the analysis of the beef case suggest a long-run relationship between the prices in both systems, though stronger in magnitude in the conventional system. Retail prices, and not farm prices, are found to respond to deviations in each system from parity in the equilibrium state. Also, data corroborates time-varying volatility. Results suggest limited volatility spilled from retail prices in both systems, but more perceptible in the conventional system from farm to retail market. Significant asymmetric effects are found in the conventional system whereas those estimated in the FQS are smaller, as could be expected *a priori* based on the targets of the European labels of origin. Further, the transmission of price volatility occurs in all directions with a strong dependence on retail prices and weak transmission from farm prices in the conventional system. In the quality system, the dependence on retail prices also affects retail and farm volatilities by increasing market volatility.

As to the lamb case study, we also find retail and farm prices to be co-integrated in both systems. This long-run relationship seems stronger in the conventional system than in the quality system but the latter presents faster adjustments in the short-run. Time-varying volatility is corroborated by the data with asymmetry. Higher and significant spill-overs are identified in the conventional system. Predicted conditional variances seem to be higher with limited fluctuations in the conventional system than in the FQS case which exhibits extremely low levels of volatility implying reduced transmission along the supply chain. Conditional variance equations indicate that conventional retail price volatility depends on its own past volatility, past farm volatility and the past correlation between these two. We also find higher sensitivity to price decreases than increases. For the FQS, price volatility is directly affected by its own past volatilities (higher in the retail market than at the farm level) but asymmetries do not seem significant.

As to the cheese case study, results reveal a long-run equilibrium relationship between the prices analysed in both the conventional and FQS case, though the relationship in the latter seems stronger and is significant. Conventional retail prices are characterized by a faster adjustment compared to conventional farm prices which are characterized by slow adjustments to deviations from the equilibrium implying retail market power. Conversely, for the FQS, farm prices are characterized by a very low adjustment and retail prices do not respond. Moreover, the results corroborate the presence of time-varying volatility along the two supply chains, conventional and FQS. The conventional system exhibits higher volatility spilled over from retail to farm level. For the FQS, the magnitude is almost negligible. In the conventional system, conditional variance equations indicate that a decrease in price instability is not guaranteed. Also, it can be seen that retail does not exert much market power. Asymmetric effects are only found to be significant for farm price volatility in the conventional system.

In sum, the results of these three case studies indicate that, in general, there is a long-run relationship whose responses in the short-run are usually faster in retail markets in FQS for meat products, whereas responses are faster in the conventional system for the cheese product. Data corroborates the existence of time-varying volatility. The magnitude of price volatility patterns is higher in conventional systems than in FQS, and usually in retail markets (beef and cheese products). And, finally, asymmetric dynamics are more significant in the conventional system, which means that, at least for these case studies, European FQS have proven to be useful in reducing price volatility linkages between chain actors.

The results derived contribute to the literature in the following aspects. They provide novel empirical results on the analysis of price linkages in levels and in volatility transmission. They focus on two different sectors (meat and cheese) of two Mediterranean countries and also compares the results obtained for the FQS designated products to those of their respective conventional counterparts. The study provides more evidence on the performance of the methodology used to account for the presence of asymmetries in the variance-covariance matrix (Kroner and Ng, 1998). Finally, this study has employed a reliable database which contains price time-series recorded systematically for two levels of the food marketing chain, farm and retail, for FQS and conventional systems.

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The Strength2Food project in a nutshell

Strength2Food is a five-year, €6.9 million project to improve the effectiveness of EU food quality schemes (FQS), public sector food procurement (PSFP) and to stimulate Short Food Supply Chains (SFSC) through research, innovation and demonstration activities. The 30-partner consortium representing 11 EU and four non-EU countries combines academic, communication, SMEs and stakeholder organisations to ensure a multi-actor approach. It will undertake case study-based quantitative research to measure economic, environmental and social impacts of FQS, PSFP and SFSC. The impact of PSFP policies on nutrition in school meals will also be assessed. Primary research will be complemented by econometric analysis of existing datasets to determine impacts of FQS and SFSC participation on farm performance, as well as understand price transmission and trade patterns. Consumer knowledge, confidence in, valuation and use of FQS labels and products will be assessed via survey, ethnographic and virtual supermarket-based research. Lessons from the research will be applied and verified in 6 pilot initiatives which bring together academic and non-academic partners. Impact will be maximised through a knowledge exchange platform, hybrid forums, educational resources and a Massive Open Online Course.

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