UNIVERSITAT POLITÈCNICA DE CATALUNYA

DEPARTAMENT D’ENGINYERIA AGROALIMENTÀRIA I BIOTECNOLOGIA

Doctoral Thesis:

“An assessment of the impacts of recent food market shocks on food prices using price transmission analysis”

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PhD Program**: Sustainability**

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Abstract

Food markets worldwide have been strongly affected by recent shocks such as food scares and the outbreak of the biofuels market. The present thesis makes a significant contribution to the existing literature on price transmission by shedding light on the impacts that these food market shocks have had on food price levels and stability. To do so, recent developments in time series econometrics are applied.

Three specific objectives have been pursued in three papers that constitute the main body of the dissertation. In the first paper, a regime-switching vector error correction model is applied to monthly price data to assess the impact of BSE outbreaks on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. Different regimes within the model represent different price behavior under different market conditions. To evaluate whether different magnitudes of the BSE food scare elicit different food price responses, a BSE food scare information index is developed and used as the variable determining regime-switching. Results suggest that BSE scares affect beef producers and retailers differently. While consumer prices are not found to respond to BSE scares, producer prices are adjusted as a response to the crisis. The magnitude of the adjustment is found to depend on the magnitude of the food scare.

In the second paper, a bivariate smooth transition vector error correction model is applied to monthly poultry price data to analyze the effects that avian influenza has had on price transmission along the Egyptian poultry marketing chain. As in the previous paper, in order to reflect consumer awareness of the crisis, an avian influenza food scare information index is developed and used within the model as a transition variable. While food scare information indices have been used to assess the economic impacts of food scares on developed countries, ours is the first attempt to use them in the context of a developing country. Our results show that price responses to deviations

from the market equilibrium parity depend on the magnitude of the avian influenza crisis. Further, these adjustments are found to have very different implications for market equilibrium: during the crisis retailers use their market power to increase marketing margins. In contrast, wholesaler margins are found to decline. Results also suggest that food safety information indices contribute to a better understanding of the economic effects of food scare crises in developing countries.

In the last paper, error correction models estimated both using multivariate local linear regression and conventional parametric techniques are applied to assess price linkages and price transmission patterns between food and energy prices in Spain. More specifically, the models study the links between biodiesel, sunflower and crude oil prices. Results suggest the existence of a long-run, equilibrium relationship between the three prices studied. Biodiesel is the only variable that adjusts to deviations from this long-run relationship. Local linear regression techniques show that the speed of adjustment of biodiesel prices is faster when biodiesel is relatively cheap than when it is expensive. Energy prices are found to influence sunflower oil prices through short-run price dynamics.

Resumen

Los mercados mundiales agroalimentarios se han visto recientemente afectados por shocks importantes tales como crisis alimentarias y el estallido del mercado de los biocombustibles. La presente tesis contribuye de forma significativa a la literatura existente sobre la transmisión de precios, analizando los impactos que estos shocks de mercado han tenido sobre los niveles de precios de los alimentos y la estabilidad de los mismos. Para ello se aplican modelos de econometría de series temporales recientemente desarrollados.

La tesis estudia tres cuestiones principales en tres artículos científicos, que constituyen el elemento central de la misma. En el primer artículo, un modelo de corrección del error de cambio de régimen se aplica a series de precios mensuales para determinar el impacto de la crisis de la EEB sobre la dinámica de los precios del bovino a lo largo de la cadena comercial en España. Diferentes regímenes dentro del modelo representan diferentes comportamientos de precios ante distintas condiciones de mercado. Para esclarecer si el efecto de la EEB sobre la transmisión vertical de precios depende de la magnitud de la crisis, se construye un índice de información sobre la EEB que se utiliza para determinar el cambio de régimen. Los resultados sugieren que la crisis de la EEB afecta a los productores de carne y a los minoristas de forma diferente. Mientras que los precios al consumo no se ajustan ante la crisis, los precios al ganadero sí que lo hacen. La magnitud del ajuste depende de la magnitud de la crisis de la EEB.

En el segundo artículo, un modelo bivariante de corrección del error de transición suave se aplica al estudio de los precios mensuales de las aves de corral para analizar los efectos que la gripe aviar ha tenido sobre la transmisión de los precios a lo largo de la cadena comercial de aves en Egipto. Al igual que en el artículo anterior, se construye un índice para reflejar el nivel de información que los consumidores tienen sobre la crisis, el cual se utiliza en el modelo como

variable de transición entre distintos regímenes. Si bien los índices de información se han utilizado ampliamente para estudiar los impactos de las crisis alimentarias en los países desarrollados, este trabajo es el primero en utilizarlos para estudiar los efectos de estas crisis en países en vías de desarrollo. Nuestros resultados sugieren que los ajustes de precios ante desequilibrios del mercado dependen de la magnitud de la crisis de la gripe aviar. Además, estos ajustes tienen consecuencias muy diferentes para el equilibrio del mercado: durante la crisis los minoristas utilizan su poder de mercado para aumentar los márgenes de comercialización. Por el contrario, los márgenes del mayorista disminuyen. Los resultados también sugieren que los índices de información sobre las crisis contribuyen a una mejor comprensión de los efectos económicos de las mismas en los países en desarrollo.

En el último artículo se analiza la transmisión de precios entre los mercados españoles de energía y de alimentos mediante un modelo de corrección del error estimado con técnicas de regresión lineal local multivariante y técnicas paramétricas convencionales. Concretamente, el modelo empírico relaciona el precio del biodiesel, el del aceite de girasol y el del crudo. Los resultados sugieren la existencia de una relación de equilibrio a largo plazo entre los tres precios estudiados. El biodiesel es la única variable que se ajusta ante desequilibrios de esta relación a largo plazo. El modelo de regresión lineal local muestra que la velocidad del ajuste de los precios del biodiesel es más rápida cuando el biodiesel es relativamente barato, que cuando es caro. Los precios de la energía también influyen sobre los precios del aceite de girasol a través de la dinámica de precios a corto plazo.

Dedication

To my parents

إلى أبي وأمي

Acknowledgments

First of all, without the guidance of ALLAH, this work would not have been done successfully.

I would like to express my sincere thanks and gratitude to my dissertation director, Teresa Serra, for her detailed supervision, support, patience, and encouragement throughout my PhD studies. You have given so much of yourself to help me succeed, I owe you so much.

I would like to thank my dissertation tutor, José Maria Gil, for his valuable advice, support and friendly and closely treatment throughout the study.

I am deeply indebted to "el Centre de Recerca en Economia i Desenvolupament Agroalimentari CREDA-UPC-IRTA" for invaluable help and for providing financial support during my PhD study.

I specially wish to express my sincere thankfulness to my parents for their continued support and encouragement for each new adventure I decide to undertake, tremendous love and many prayers that have been always very valuable to me.

I would also like to express my deep thanks to my wife for her patience and support in both the easy and hard times that I have faced.

My deep thanks also to my brother-in-law, Allam, for his support and encouragement during my study.

I cannot conclude without acknowledging the continuous and endless support of my brothers, Kazem and Ehab, and my sisters and all my family.

Finally, I would like to express my gratefulness to my friends Amr, Faical, Ismail, Talha, Youssef, Mohammed, Musbah, Bouali, Ahmed, Fadi. Without you guys, I would not have made it this far.

Chapter 1

Introduction

Food prices in the 2000s decade were characterized by higher instability relative to previous decades (FAO-OECD, 2011; FAO, 2008). Different underlying reasons providing an explanation to this price instability have been identified and include, among others, important food scares such as the Bovine Spongiform Encephalopathy (BSE) and the Avian Influenza (AI), and the outbreak of the biofuels market that have involved important changes in food markets. As emphasized by FAO (2008), late changes in agricultural prices have not been uniformly passed along food marketing chains. As a general rule, consumer prices have experienced considerably smaller changes than producer prices. This dissertation focuses on shedding light on this issue by studying the impacts that the market shocks mentioned above have had on food price behavior.

Recent worldwide food scares have raised public awareness on food safety and substantially altered food demand. Apart from the obvious concerns about the impacts of these scares on human and animal health, there are also worries about the effects that these crises have had on price transmission along the food marketing chain. Suspicions have arisen that food scares have had a different impact on retailers and producers. It has been argued that agricultural producers usually suffer stronger price declines than retailers, increasing the marketing margins in favour of the latter (Lloyd et al., 2006) and threatening farmers' standard of living.

The rapid growth in worldwide biofuel production over the last decade, which mainly relies on agricultural feedstocks, has also generated social and political concerns with regards to the impacts that biofuels can have on food markets. Previous research has shown that the outbreak of the biofuels market has strengthened the links between food and energy prices (Balcombe and Rapsomanikis, 2008; Zhang et al., 2009b; Serra et al., 2011a and 2011b). This has brought the question of whether biofuels contribute to increased food prices.

The objective of this thesis is to shed light on the impacts of recent food market shocks on food price behavior. To do so, we apply recent developments in time series econometrics that allow characterizing price transmission along the marketing chain. Time series econometrics have been widely used in the economics literature to assess price behavior in agrofood markets. This literature has paid special attention to understand the way in which price changes at one level of the market are transmitted to other market levels.

Vertical price transmission allows an approximation to the overall operation of the market (Goodwin and Holt, 1999). Latest developments in time series econometrics that involve substantial improvement in price modeling relative to older proposals, have contributed to renewed interest in price transmission analyses. Also, recent shocks affecting food markets have raised concerns about the extent and speed by which these shocks are passed along the market chain. This has implied further attention devoted to this literature. Price transmission analyses based on time series models have the advantage that, in being a reduced form of structural models, they only require price data. As a result, an in contrast to structural models, their estimation is not restricted by data availability.

Many empirical analyses have found evidence of non-linear (asymmetric) vertical price transmission characterizing several different markets (Peltzman, 2000). A number of institutional and theoretical explanations have been suggested as a justification for nonlinearities. Market power, search costs, adjustment costs, among others, can lead to nonlinear price behaviour (see, for example, Lloyd et al., 2006; Defra, 2004; Reagan and Weitzman 1982, or Azzam 1999). In assessing the impacts of market shocks on price behaviour, this dissertation uses flexible models that allow for these nonlinearities.

Different econometric techniques have been developed to model non-linear price adjustments. Early analyses based on the econometric model introduced by Wolffram (1971) and later refined by Houck (1977) and Ward (1982) have been criticized for ignoring the time-series properties of the (usually non-stationary) data. Tong (1978) originally introduced nonlinear threshold time series models. Tsay (1989) developed a methodology to test for threshold nonlinearity based on autoregressions and to model threshold autoregressive (TAR) processes. Balke and Fomby (1997) combined error correction frameworks and the TAR model developed by Tsay (1989) into a Threshold Vector Error Correction Model (TVECM).

TVECMs have been criticized because they rely upon the assumption that transition between different price behavior regimes is abrupt and discontinuous. Teräsvirta (1994) developed smooth transition models that are less restrictive in that they allow for smooth shifting between regimes and contain discontinuous shifting as a special case. Smooth transition models have been recently generalized and applied in the context of multivariate VECM (see van Dijk et al., 2002). A further generalization of smooth transition models can be achieved by implementing local polynomial regression (non-parametric) techniques (see Fan and Gijbels, 1996). The objective of my thesis is to implement these latest advances in time-series econometrics to assess the impacts of the recent food market shocks on price behaviour in agrofood markets.

In particular, three specific objectives have been defined:

a. First, to assess the impacts of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain.

b. Second, to study the effects of the AI food scare on poultry price behavior in a developing country: Egypt.

c. Third, to analyze price linkages and price transmission patterns between biodiesel, sunflower and crude oil prices in Spain.

To achieve the first objective, a regime-switching vector error correction model (RSVECM) is estimated. Different regimes within the RSVECM represent different price behavior under different market conditions. The degree to which price transmission is affected by a food scare crisis is likely to depend on the scale of such crisis. To capture this issue, a BSE Food Scare Information Index (FSII) is built and used as the variable determining regime-switching. Hence, it is implicitly assumed that different levels of the food scare can lead to different price responses and that the food scare index leads to switching between different price behavior regimes.

It is worth noting that, in spite of previous attempts to characterize price transmission responses to food scares, no previous analysis has allowed for regime-switching linked to the magnitude of the crisis, which represents a contribution of this dissertation to the literature. While previous literature has shown that the BSE crisis had an impact on vertical price transmission processes within the UK beef market (Lloyd *et al.*, 2001), this dissertation is the first to focus on the impacts of BSE on price transmission within the Spanish beef market.

The second objective of this thesis is to investigate the effects of the AI on price transmission along the Egyptian poultry marketing chain. According to the World Health Organization (WHO), Egypt has been the third most affected country by this crisis all over the world, after Indonesia and Vietnam. In order to achieve the second objective, a bivariate Smooth Transition Vector Error Correction Model (STVECM) is estimated so as to capture the long-run relationship between wholesale and consumer prices, the (possibly non-linear) adjustments toward long-run parity, as well as the short-run dynamics. Like RSVECMs, STVECMs are able to distinguish between different regimes that represent different price behaviour under different economic conditions. In contrast to RSVECMs that assume that transition between different price-behaviour regimes occurs abruptly and discontinuously, STVECMs allow for a smooth transition. To our knowledge, this is the first research that assesses the impacts of a food scare by using a STVECM. Also, our analysis contributes to previous literature in that it focuses on studying the effects of AI on price transmission along the Egyptian poultry marketing chain, a market that has not been investigated yet, although, as we have mentioned, Egypt has been one of the most affected countries.

Recent increases in biofuel production pursuing different objectives such as to reduce the dependence on crude oil, diversify energy supplies, support rural economies or reduce greenhouse gas emissions, have raised questions regarding the links between energy and food price levels. Currently, food inputs represent the main feedstocks used in biofuel production. In Spain, for example, biodiesel production is mainly based on sunflower oil. The demand for sunflower oil to produce biodiesel directly competes with its use to produce food or animal feed. It is thus important to determine the extent to which energy (biodiesel) markets have the potential to increase food (sunflower oil) prices. The third objective of this thesis is to shed light on this issue.

While previous literature has focused on studying price linkages and price transmission patterns between energy and feedstock markets in the US and Brazil, the two countries that dominate the international ethanol market (Balcombe and Rapsomanikis, 2008; Zhang et al., 2009a; Serra et al., 2011a and 2011b), European markets, that represent about 65% of global biodiesel output, have not yet received any attention. This dissertation fills this gap in the literature by assessing the relationship between biodiesel, sunflower and crude oil prices in Spain, the third largest biodiesel producer within Europe.

Here, unlike the previous parts of the thesis and also unlike the existing price transmission literature, which reveals that the analyses of the dynamic price relationships have been typically based on parametric models, we are interested in studying the relationship between energy prices and the sunflower oil price in Spain by using multivariate local polynomial regression (non-parametric) techniques. Non-parametric techniques are data-driven methods that do not require any assumptions about the functional form characterizing price behavior and thus are robust to misspecification issues. The results of non-parametric techniques are compared with the ones derived from the parametric error correction model.

In addition to this introductory chapter and a concluding remarks section at the end of the thesis, the dissertation is organized into three main chapters addressing the specific objectives discussed above. These chapters have been written as a scientific article and have been submitted for publication consideration in different scientific journals. The first paper, chapter 2, studies the effects of the BSE crisis on price relationships and patterns of transmission among farm and retail markets for bovine in Spain. The paper has been published in *Agricultural Economics* (Hassouneh et al., 2010). The second paper, chapter 3, addresses the second objective of the study, which is the analysis of the impacts of the AI on price transmission along the Egyptian poultry marketing chain. The article is currently being considered for publication in the *Food Policy* journal (second round review). The analysis of the relationships between biodiesel, sunflower and crude oil prices in Spain is presented in the third article in chapter 4. The article is presently being considered in the *Energy Economics* journal.

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Chapter 2

Price transmission in the Spanish bovine sector: the BSE effect

1. Introduction

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

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

Second, the Spanish beef sector has substantial economic and social relevance both within Spain and within the EU. Beef is one of the most important activities in the Spanish agricultural sector, representing almost a fifth of the Spanish gross animal production (Ministerio de Medio Ambiente y Medio Rural y Marino, 2006). Also, Spain occupies the fifth position in the EU rankings for bovine production with 671 thousand tons, representing 9% of total output (Eurostat, 2007). The Spanish beef sector has relevance from a social perspective as well, since it constitutes, together with sheep

[1] EU refers to EU-15.

farming, the main economic activity that can be carried out in several less-favored mountain areas. It is thus interesting to see how the BSE crisis affected the overall operation of this important economic sector.

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

2. The BSE in Spain

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

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

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

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

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

3. Previous research: a literature review

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

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

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

Another common finding of previous analyses is that prices at the retail level are sticky or slowly-responsive to price changes at other levels of the marketing chain.

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

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

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

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

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

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

4. Econometric methods

4.1. Regime-Switching Autoregressive Models

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

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

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

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

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

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

4.2. Regime-Switching Vector Error Correction Models

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

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

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

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

where:

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

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

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

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

The parameters of the multivariate RSVECM can be estimated using sequential multivariate least squares in two steps. In the first step, a grid search is carried out to estimate the threshold parameter, c.[4] The threshold is searched over the values of the threshold variable and the search is restricted to ensure an adequate number of observations for estimating the parameters in each regime. Recently, Serra and

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

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

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

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

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

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

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

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

4.3. Specification tests

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

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

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

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

4.4. The food scare information index

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

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

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

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

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

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

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

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

5. Empirical application

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

As noted above, health risks have received increasing attention among consumers in developed countries and strengthened the relationship between food scares, food consumption, and prices. More specifically, if consumers believe that beef is unsafe to eat, there will be a decline in the demand and possibly the price of beef. To investigate the impacts of food scare concerns during the BSE crisis on price

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

relationships and patterns of vertical price transmission within the Spanish bovine sector and following the methodology outlined in subsection 4.4, we construct a FSII that captures the degree of the consumer food scare. The index is based on a monthly count of newspaper articles on the BSE crisis appearing in a major Spanish newspaper, EL PAÍS.

Weinberger and Dillon (1980) suggest that supporting articles (unfavorable news) may be more influential than a similar amount of non-supporting articles (favorable news). It has been also reported that a similar quantity of unfavorable news weighs far more heavily on consumer decision-making than favorable news (Chang and Kinnucan, 1991; Kinnucan and Myrland, 2000). However, Mazzochi (2006) argues that discrimination between favorable and unfavorable articles is highly subjective. This can be especially true with the BSE crisis, since the long latent period of vCJD will not cause the same impact on the young than on the old population. Due to the aforementioned reasons and following Kim and Chern (1997) and Chern and Zuo (1997), we do not weigh articles depending on whether they are favorable or unfavorable.

We construct our monthly FSII over the period from July 1995 to December 2005, by scanning all articles relevant to BSE crisis using different keywords.[6] The keywords searched were “Vaca(s) Loca(s)” (mad cow), “Encefalopatía Espongiforme Bovina” (Bovine Spongiform Encephalopathy), and “Creutzfeldt-Jakob.” The number of news articles ranged from a maximum of 354 in February 2001, to a minimum of 0 in a few months far from the peak of the crisis. The average was 20.8 news articles per

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

month with a standard deviation of 44.12. By excluding geographical indicators from the keyword list, our FSII intends to include both the information on the development of the crisis in Spain and in other EU countries. Due to the important trade relationships between Spain and the EU, Spanish markets are likely to have been affected by the development of the food scare in other EU countries. As a result, despite the fact that the BSE was first identified in November 2000 in Spain, the FSII is developed from 1995, in order to capture the impacts of earlier infections in other countries.

Following Chern and Zuo (1997), we computed our monthly FSII using different values for both the number of lags (n) and peak times (m) and found the index to remain relatively stable across these values. In light of these results, we selected n=6 which is consistent with the recommendation by Clarke (1976) and requires m to be equal or less than two. Since the FSII is very similar independently of the chosen value for the peak time, we selected m=1.[7]

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

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

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

6. Results

As noted above, our empirical analysis utilizes monthly farm-gate and consumer beef prices. It is also based upon a newspaper index that is a measure of the degree of food scare on a monthly basis. The empirical analysis is based on a logarithmic transformation of prices. Standard Dickey and Fuller (1979) and KPSS (Kwiatkowski *et al.*, 1992) tests for each price series provide evidence that all price series are integrated of order one I(1).[8]

After testing for unit roots, the Engle and Granger (1987) test for cointegration is used. In doing so, the equilibrium relationship is normalized by the producer price and ordinary least squares (OLS) are used to obtain estimates of the cointegrating parameters (see table 1). The normalization variable is selected according to previous research results that confirm that while producer prices tend to adjust to their long-run parity, consumer prices are sticky or slowly-responsive to price changes occurring at other levels of the marketing chain (see Goodwin and Holt, 1999; Borenstein *et al.*,

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

1997; Peltzman, 2000). This is confirmed later in our analysis. The Engle and Granger (1987) test indicates the existence of a cointegration relationship among producer and consumer price series (see table 2).[9] Other analyses have also found evidence of cointegration between producer and consumer prices for beef (Lloyd *et al.*, 2001; Sanjuán and Dawson, 2003; Jaenicke and Reiter, 2003; Livanis and Moss, 2005).

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

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

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

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

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

FSII>31.218. In table 4 we present the RSVECM parameter estimates across the different regimes.

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

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

While our results are compatible with previous research results on the BSE crisis, the use of a RSVECM offers an advantage over previous studies based on

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

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

7. Concluding remarks

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

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

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

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

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

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

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

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Figure 1. Evolution of the number of BSE cases in Spain (2000-2007)

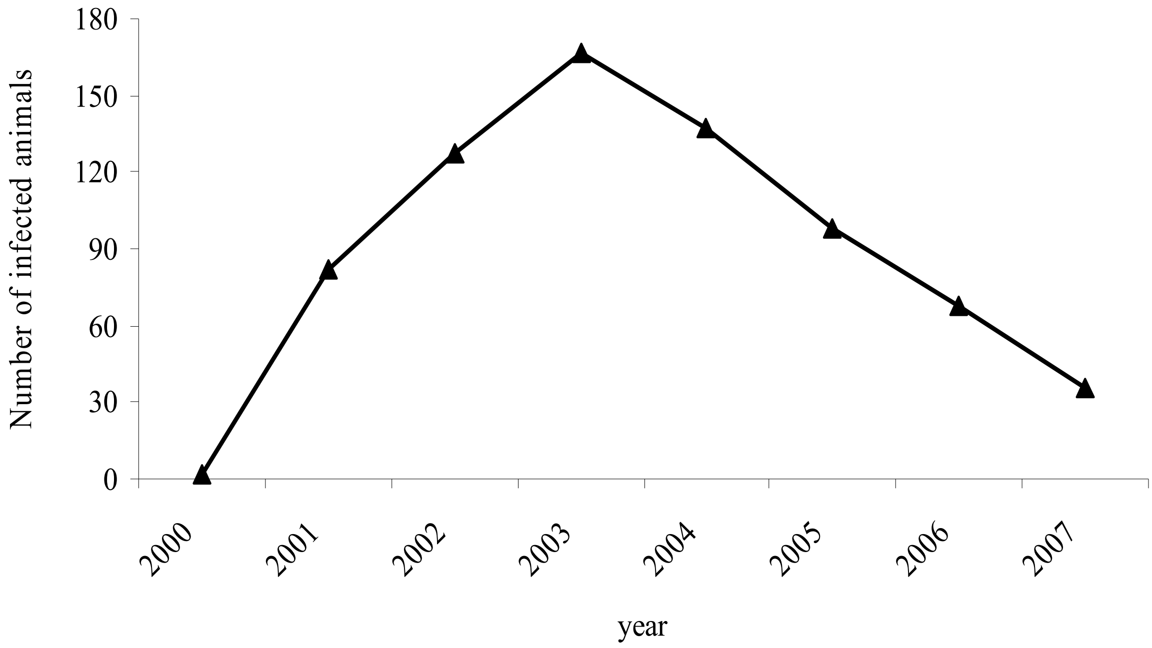

Figure 2. Food scare Information index.

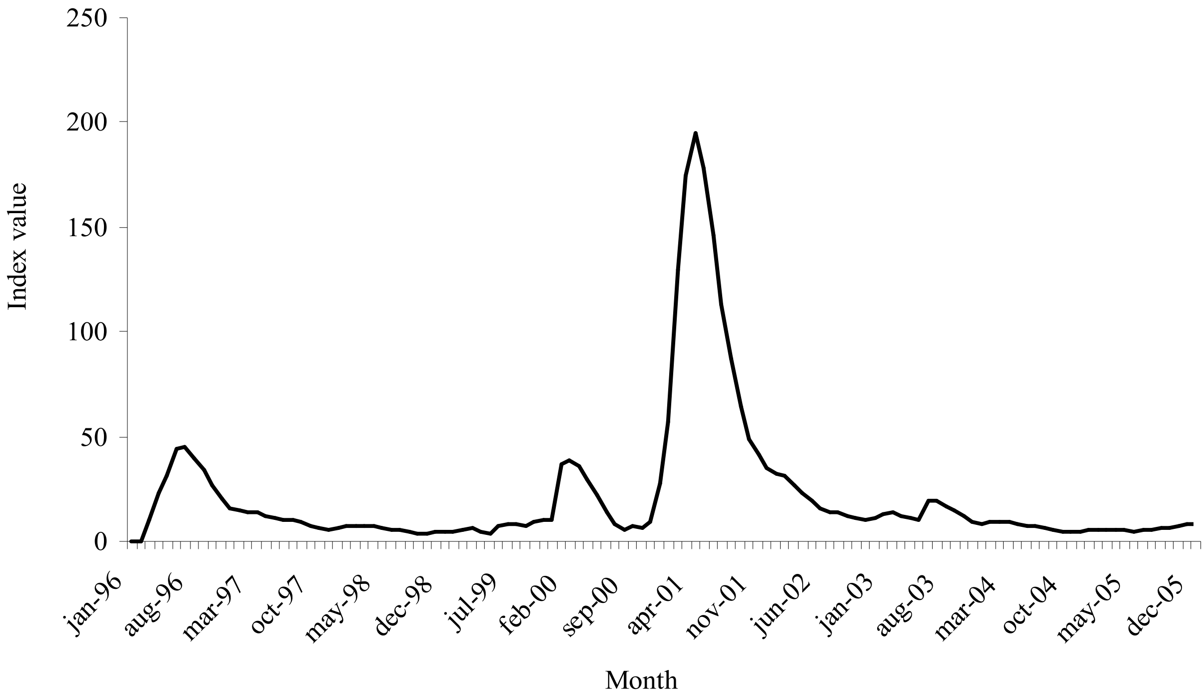

Figure 3. Timing of jumps among regimes

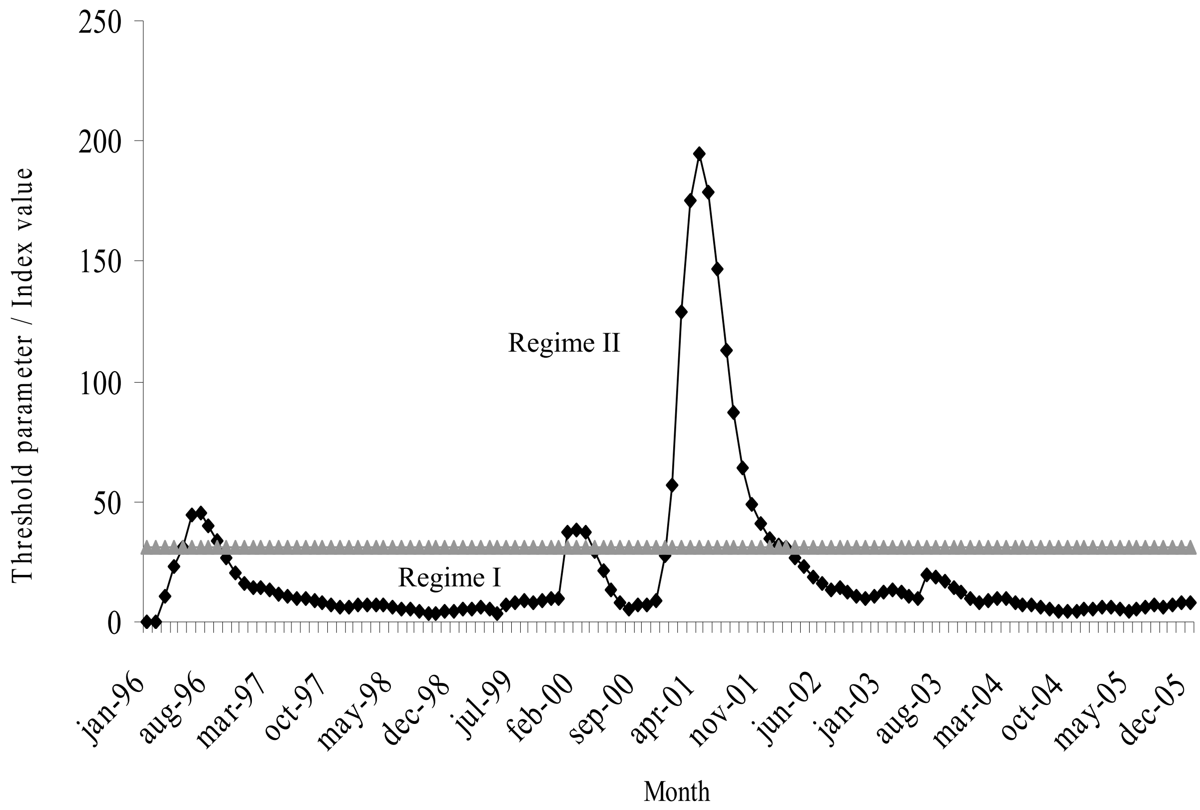

Table 1. OLS Estimates of the cointegrating relationship

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

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

Table 2. Engle and Granger test for cointegration

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

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

Table 3. RSVECM: Thresholds and Sup-LR test

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

Table 4. RSVECM: parameter estimates

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

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

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

Chapter 3

Food scare crises and developing countries: the impact of avian influenza on vertical price transmission in the Egyptian poultry sector

1. Introduction

Food safety has become a key issue for consumers and, consequently, a relevant objective of food policy in many countries. The food scares that have been appearing over the last few years (for example dioxins, Bovine Spongiform Encephalopathy (BSE), Foot and Mouth Disease (FMD), Avian Influenza (AI)) have contributed to increased public awareness of food safety. The globalisation of markets and the increasing speed of and access to information have amplified these concerns to a global scale. As consumers' sensitivity towards food safety and production standards has increased, they have changed their consumption habits. While this once was a phenomenon confined to developed countries, concerns over food safety have now spread to developing countries too, which often find addressing such issues even more problematic.

The AI outbreak originated in Southeast Asia and quickly spread worldwide. Defence mechanisms against food scares implemented in developed countries, mainly as a result of previous Mad Cow and the Foot and Mouth disease episodes, helped to mitigate the impact of AI in these countries. However this was not the case in countries like Egypt which, with the advent of AI, faced a nationwide food scare for the first time.[1] The crisis dramatically affected the Egyptian animal production sector and, more specifically, the poultry sector that suffered the culling of more than 40 million birds. The poultry sector in Egypt is very relevant in that it represents the main and cheapest

[1] While Egypt had been previously affected by other animal disease crises, they were all confined to particular regions of the country and did not gain massive mass media attention as it was the case with AI.

source of animal protein for Egyptian consumers who have average animal protein intakes quite below the world average (Hosny, 2006).

Apart from the obvious impact that AI has had on human health, there are also worries about the economic effects of the crisis. This paper is the first to assess the effects of AI on price transmission along the food marketing chain within the Egyptian poultry sector. The analysis by Saghaian et al. (2008) was the first to investigate the impact of AI on price transmission within the Turkish food marketing chain. These authors provided evidence that the AI crisis had had an impact on vertical price transmission processes. Despite their attempt to characterize price transmission responses to the AI food scare, they did not develop a Food Scare Information Index (FSII) to reflect consumer awareness of the AI crisis. While food scare information indices based upon a count of newspaper articles have been widely used to assess the economic effects of food contamination outbreaks in developed countries, ours is the first study that attempts to determine the impact of news on price behaviour in developing economies.

Due to high illiteracy levels (29.7% in Egypt in 2006 (CAPMAS, 2006)) and low income levels precluding many people to buy newspapers, these indices may not be regarded as representative measures of developing countries' public awareness of the AI crisis.[2] The indices, however, are indeed a good approximation to consumers' awareness because of the high correlation between news in different information channels (TV, newspapers, mosques, etc.). This correlation is expected to be higher in a strictly regulated country such as Egypt, where all information channels are controlled by the

[2] While TV could be considered as an alternative valid information source on the food scare crisis, records on the news published in this media are not publicly available.

Egyptian Ministry of Mass Media. This results in newspapers and other mass media sending very similar messages to the audience. Our study is the first attempt to test the validity of this type of index to assess the economic impacts of food scares in developing countries, thus shedding light on this interesting discussion.

To achieve our objective, we have estimated a bivariate Smooth Transition Vector Error Correction Model (STVECM) that enables capturing the long-run relationships between the variables of the model, the (possibly non-linear) adjustments toward long-run parity, as well as short-run dynamics. STVECMs are able to distinguish between different regimes that represent price behaviour under different economic conditions. The FSII is used as the variable determining regime-switching, which allows us to investigate whether prices within the supply chain respond differently to distinct levels of food scares. To our knowledge, this is the first research that assesses the impacts of a food scare by using a STVECM.

The rest of the paper is organised as follows. Section 2 presents an overview of the Egyptian poultry sector and the AI crisis. A literature review is carried out in section 3. The econometric methods are explained in section 4. The data used, the construction of the information index and the results are described in section 5. The paper ends with the concluding remarks section.

2. Egyptian poultry sector and the AI crisis

2.1. Egyptian poultry sector

The poultry sector is the second most important animal protein provider after the bovine sector in Egypt. Between 2003 and 2006 investment within the poultry sector reached

20 billion LE.[3] The sector employed about 2.5 million workers in permanent and casual jobs and produced 95 million chickens per year. Gross production in 2005 was on the border of 10 billion LE, which represents around 20% (7%) of total animal (agricultural) production (Ministry of Agriculture, 2006). Poultry production systems in Egypt are quite diverse, ranging from rural, very small-scale, free-range poultry production to highly intensive caged systems. Poultry production has been one of the fastest growing industries in Egypt during the1990s with an annual growth rate of 8.7% (Taha, 2003).

Hosney (2006) indicates that the founding of the National General Poultry Company (G.P.C.) marked the beginning of the modern poultry industry in Egypt. This company aims both to guarantee high quality and inexpensive animal protein for the growing Egyptian population, and to industrialise poultry production by means of adopting modern technologies and skilled management techniques.

The Egyptian poultry sector has traditionally benefitted from government intervention in the form of the provision of subsidized feed ingredients, low-rate subsidized loans from the National Agricultural and Development Bank, or the allowance of a ten year tax exemption period followed by low taxation rates. The industry has also benefitted from permissive measures such as the relaxation of regulations setting forth minimum distance requirements between different poultry farms, which made the implementation of bio security and disease controls nearly impossible. Furthermore, the government ensured high intervention levels for the Egyptian poultry industry by imposing strong restrictions on foreign poultry imports (including high tariffs or import bans). Subsidisation and protectionist policies resulted

[3] 1 US$ = 5.82 LE.

in a very fragile sector with many inefficient farms housing hidden weaknesses. The AI crisis brought to light those weaknesses and forced many small and inefficient farms to exit the sector.

Before the AI crisis consumers usually purchased live chickens that the retailer or they themselves slaughtered. This tradition made it possible to eat meat throughout the day without the need to refrigerate it. However, this practice contributed to spreading the pandemic in Egypt due to the direct contact with live animals.

2.2. AI crisis

More than 200 million birds have died or been culled worldwide as a consequence of the AI subtype H5N1 since 2003 (FAO, 2006). According to the World Health Organization (WHO, 2006) “AI is an infectious disease of birds caused by type A strains of the influenza virus. The disease occurs worldwide. While all birds are thought to be susceptible to infection with AI viruses, many wild bird species carry these viruses with no apparent signs of harm”. According to the WHO, up until February 17, 2010, there were 476 confirmed human cases of AI worldwide, of which 283 died; a fatality rate of 59%.

The first case of AI in Egypt took place in February 2006. From this date, 97 cases of infection in humans have been reported, out of which 27 were fatal. The WHO states that Egypt is the third most affected country by this crisis in the world, after Indonesia and Vietnam.

The consequences on the poultry sector have been evident. Chicken meat demand was reduced substantially, which generated a price reduction. Afterwards the market recovered, but there is still a common feeling of market fragility. Approximate estimations of the losses owing to AI in Egypt are above 2 billion LE, representing almost 25% of the production value (Ibrahim, 2006).

The Government policy instruments to deal with this crisis mainly comprise a media campaign through all possible communication channels aiming at providing advice to the population on how to deal with the food scare (Hosny, 2006) , the closure of live bird shops and the elimination of tariffs on imported poultry meat to compensate the production shortage.

3. Literature review

Understanding vertical price transmission allows for an approximation of the overall operation of the market (Goodwin and Holt, 1999). The recent development of econometric methods has greatly contributed to a better understanding of the functioning of market operations. A common finding of previous analyses is that different levels of the supply chain respond differently to distinct market shocks. In particular, upstream prices in the marketing chain are generally found to do all the adjustment, while consumer prices are sticky and slow to respond (Peltzman, 2000).

Within this context, recent studies have found that relevant food scares lead to different price adjustments within the supply chain. Lloyd et al. (2001) and Hassouneh et al. (2010) analyze the impact of the BSE outbreaks on vertical price transmission within the UK and Spain beef markets, respectively. Lloyd et al. (2001) find that information shocks are fairly transient in retail prices, but persist at the wholesale and farm levels. The analysis by Hassouneh et al. (2010) also suggests that BSE scares affect beef producers more profoundly than retailers. Serra (2011) assesses the impacts of BSE on price volatility transmission along the Spanish beef marketing chain and finds volatility links between producer and consumer prices to be altered as a result of the crisis.

Although the economic impact that the AI crisis has had on the poultry and animal production sectors in many countries worldwide has been substantial, only a few studies have tried to quantify this impact. Most of these analyses have focused on the effects of the AI outbreak on consumer behaviour and willingness to pay and on public policy decision-making (see, for example, Akben et al. 2008). Beach et al. (2008) utilize scanner data to study the effect of AI news on consumer purchasing behaviour in Italy and find a significant impact but within a limited duration. The analysis by Brown et al. (2007) investigates the impact of the AI outbreak on different agricultural sectors within the US market, finding that several different agricultural sectors were affected.

The aforementioned literature has paid little attention to the impact of AI on price transmission along the food marketing chain in developing countries. However, research conducted by Saghaian et al. (2008) is an exception: they analyze the effect of the AI crisis on producer and consumer prices within the Turkish poultry sector, providing evidence that the AI outbreak has had an impact on vertical price transmission processes.

Our paper makes use of a non-linear bivariate STVECM to assess the effects of AI on price transmission along the Egyptian poultry marketing chain.[4] We consider the

[4] Although structural models are more suited than time-series models in identifying the reasons underlying a certain form of revealed price behavior, characterizing the nature of price transmission among different market levels, which is the task accomplished by this article, is an important research agenda. Hence, when using time-series models, interpretation of research results should be based on

bivariate STVECM a suitable methodology to assess price behaviour because it allows for nonlinearities in price adjustment, different price behaviour regimes depending on the predominant economic conditions, as well as for smooth transition of prices between these different regimes. The STVECM allows for assessment not only on how individual prices adjust to their long-run equilibrium parity, but also on short-run price dynamics. Further, it permits distinguishing between different price-behaviour regimes, depending on the magnitude of the food scare crisis.

Previous literature has found that different degrees of food scares lead to different price adjustments within the supply chain. Hassouneh et al. (2010) use a Regime Switching Vector Error Correction Model (RSVECM) to capture the impacts of the BSE crisis on the adjustment of producer and retailer prices in the Spanish beef marketing chain. Their findings provide evidence that while consumer prices do not adjust to deviations caused by the crisis, producer prices are endogenous and do all the adjustments. These adjustments depend on the magnitude of the BSE crisis. The interesting analysis by Hassouneh et al. (2010) suffers from an important limitation: in using a RSVECM it assumes that transition between different price-behaviour regimes occurs in a discontinuous and abrupt way.[5] Smooth transition models used in our analysis are less restrictive in that they allow for smooth shifting between regimes and contain discontinuous shifting as a special case. By developing an AI food scare index

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profound knowledge of the market being analyzed, its structure, regulation, and other important characteristics.

[5] Serra et al. (2006) utilize both non-parametric regression and non-linear threshold models in order to study price transmission processes within EU pork markets after the implementation of the EU single market in 1993. They find that price transmission from one regime to another is likely to occur in a smooth way.

that is used as a transition variable within the non-linear STVECM, we can analyze to what extent different levels of the crisis can affect price behaviour along the Egyptian food marketing chain.

4. Methods: Smooth transition vector error correction models and food information indices

4.1. Smooth transition vector error correction models

Recent developments of time series analysis have greatly contributed to a better understanding of the dynamic behaviour of economic variables which can be observed in the real world. Threshold models originally introduced by Tong (1982) are one of the most relevant families of non-linear time series models and have attracted the attention of many previous empirical studies (see, Chavas and Mehta, 2004; Goodwin and Piggott, 2001; Serra and Goodwin, 2004; Hassouneh et al., 2010; Serra et al., 2006). Threshold models allow for recognition of nonlinear price adjustments by considering different price behaviour regimes that represent different price adjustment, depending on the prevailing economic situation. These models have been criticized since transition between regimes takes place in a discontinuous and abrupt fashion.

Smooth-transition type of models such as multivariate STVECM originally developed by Teräsvirta (1994) allow for transition to occur in a smooth fashion. STVECM also assess both the (possibly non-linear) price adjustments toward long-run equilibrium and the short-run price dynamics. Following van Dijk et al. (2002) a two-dimensional STVECM can be written as

$$
\begin{aligned}
\Delta \mathbf{P}_t = & \left(\boldsymbol{\mu}_1 + \alpha_1 \mathbf{v}_{t-1} + \sum_{j=1}^{p-1} \Phi_{1,j} \Delta \mathbf{P}_{t-j} + \lambda_1 \Delta I_{t-j} \right) (1 - G(s_{t-d}; \gamma, c)) + \\
& \left(\boldsymbol{\mu}_2 + \alpha_2 \mathbf{v}_{t-1} + \sum_{j=1}^{p-1} \Phi_{2,j} \Delta \mathbf{P}_{t-j} + \lambda_2 \Delta I_{t-j} \right) (G(s_{t-d}; \gamma, c)) + \boldsymbol{\varepsilon}_t
\end{aligned}
\tag{1}
$$

where $\mathbf{P}_t = (p_w, p_c)$ is a (2 x 1) vector of non-stationary prices containing wholesale, p_w, and consumer, p_c, prices. I is an FSII, $\boldsymbol{\mu}_i$ i=(1,2) are (2 x 1) vectors of constant terms, α_i are (2 x 1) parameter matrices representing the speed of adjustment of each price to deviations from the long-run relationship, $\mathbf{v}_{t-1} = \boldsymbol{\beta}' P_{t-1}$ is a vector containing deviations from the long-run equilibrium relationship (error correction term), β (2 x 1) contains the parameters of the cointegration relationship, $\Phi_{i,j}$, j=1,..., p-1 are (2 x 2) parameter matrices representing short-run price dynamics and $\lambda_{i,j}$ measure the short-run impacts of food scares on price behaviour. Details on how the FSII variable is developed are provided below. $\boldsymbol{\varepsilon}_t$ is a 2-dimensional vector white noise process with a mean zero vector and a (2 x 2) Σ covariance matrix.

$G(s_{t-d}; \gamma, c)$ is the smooth transition function which is assumed to be continuous and bounded between zero and one. The transition function, which will be described in more detail below, depends on the transition variable s_{t-d}, as well as on the speed of transition and threshold parameters, γ and c respectively. The STVECM can be considered a regime-switching model that allows for two regimes associated with the extreme values of the transition function, $G(s_{t-d}; \gamma, c) = 0$ and $G(s_{t-d}; \gamma, c) = 1$. The transition from one regime to the other takes place in a smooth way. The transition function is specified using an exponential functional form (Teräsvirta, 1994).[6]

The exponential transition function can be expressed as follows:

$$G(s_{t-d};\gamma,c)=1-\exp\left(-\frac{\gamma(s_{t-d}-c)^2}{\sigma^2(s_{t-d})}\right),\ \gamma>0 \tag{2}$$

After combining (1) and (2) we obtain an exponential STVECM (ESTVECM). Parameter γ determines the speed of the transition from one regime to the other, c is the threshold between the two regimes, s_{t-d}, as noted, is the transition variable and $\sigma(s_{t-d})$ is the sample standard deviation of s_{t-d} used as a normalization variable. Under this approach, the adjustment is symmetric around the parameter c, but differs for large and small absolute values of s_{t-d}. In other words, $G(s_{t-d};\gamma,c)$ takes values close to zero for values of s_{t-d} near c, and when s_{t-d} takes relatively large values with respect to c, the transition function will be close to 1.

Since our paper focuses on the effects that AI had on price transmission along the Egyptian poultry marketing chain, a food scare index, based on the number of news articles published on the AI crisis, is developed and used as the transition variable s_{t-d}. In setting s_{t-d} equal to the food scare index, we assume that different levels of food scare can lead to different values of the transition function and different price behaviour

[6] While it is also usual to specify this function using a logistic form, model specification tests supported the use of an exponential function.

regimes.[7] A priori we expect the STVECM will allow for distinguishing between two extreme price behaviour regimes: one characterized by relatively calm markets and the other characterized by tumultuous periods.

Teräsvirta (1994) suggests three important steps in order to estimate a STVECM. First, a linear vector error correction is estimated in order to determine the optimum lag of the predictors of the model. To do so, the AIC and SBC criteria are used. A test for linearity against a STVECM is then applied.

A test for linearity in equation (1) is equivalent to testing whether the following null hypothesis can be accepted: $H_0 : \gamma = 0$. If this happens, the non-linear STVECM reduces to a linear VECM. A problem related to testing $H_0 : \gamma = 0$ against $H_1 : \gamma \neq 0$ is that equation (1) is not indentified under the null hypothesis. We follow Luukkonen et al. (1988) who solve this problem by replacing the transition function with a suitable Taylor series expansion. A system likelihood ratio statistic or an F-test can then be used in order to test for linearity. Finally, we select the optimal functional form for the transition function following the sequence of nested tests proposed by Teräsvirta (1994) and Luukkonen et al. (1988).

4.2. Food scare information index

Several methods have been introduced in order to construct a FSII based on a news count. More specifically, Smith et al. (1988) define the index as the actual number of articles published on the topic of interest in each period. Brown and Schrader (1990) propose a different technique, which does not ignore previously published articles, to

[7] Hassouneh et al. (2010) utilize a FSII as a threshold variable within a RSVECM and show that the degree to which price transmission is affected by a food scare crisis depends on the scale of the crisis.

construct a cholesterol information index for their study of shell egg consumption in the U.S. They developed their index by accumulating the number of supporting articles (unfavourable news) minus the non-supporting articles (favourable news) using equal weights for these two types of articles. Chern and Zuo (1997) extend the cumulative method used by Brown and Schrader (1990) by building a new fat and cholesterol information index that considers a carryover weight for articles. The articles are also assumed to have a finite duration and lag distribution as a source of information. In our study we follow this method.[8]

The FSII based on Chern and Zuo (1997) can be expressed as:

$$FSII_t = \sum_{i=0}^{n} W_i \ \mathrm{NM}_{t\text{-}i} \tag{3}$$

where NM_{t-i} is the number of relevant articles (both supporting and non-supporting) published during period *t-i*, W_i is the weight attributed to the lagged period, and *n* is number of lagged periods considered. This method not only allows for a carryover effect but also for a decay effect of news as a source of information. The carryover and decay effects are captured by specifying the weight function and the total lag period. Early literature has used the second order polynomial function for constructing food scare indices (Ward and Dixon, 1989). This function has been criticized since it generates symmetric weights (see Chern and Zuo, 1997 for further detail). To overcome this problem, Chern and Zuo (1997) propose the Cubic Weight Function (CWF):

[8] This particular FSII method has been used by Hassouneh et al. (2010), among others.

$$W_i = \delta_0 + \delta_1 i + \delta_2 i^2 + \delta_3 i^3 \tag{4}$$

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

$$W_i = 2a/((n+1)b) + (12m/b)i - (6(n+1+m)/((n+1)b))i^2 + (4/((n+1)b))i^3 \tag{5}$$

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

5. Results

Our empirical model utilizes two series of monthly wholesale and consumer poultry prices obtained from the Egyptian Central Agency for Public Mobilization and Statistics

(CAPMAS, 2008).[9] The average wholesale price equals 6.63 LE per kilogram with a standard deviation of 1.44, while the average retail price equals 7.20 LE per kilogram with a standard deviation of 1.38 (see Figure 1). Our data set extends from January 2003 until December 2006, with a total number of 48 observations.[10]

In order to approximate consumer awareness of the AI crisis through a FSII, we carried out a monthly count of newspaper articles published in the most popular Egyptian newspaper, Alahram, from January 2003 to December 2006. We searched for those articles containing the keyword "avian influenza". The average number of published news pieces is 29 per month with a standard deviation of 40. The maximum number of news articles was 214 in March 2006, following confirmation of the first Egyptian AI case in February 2006. The minimum number of news articles is zero.

No discrimination between positive or negative messages (as in Verbeke and Ward, 2001; Liu et al., 1998; and Smith et al., 1988, among others) is carried out because, as indicated by Mazzocchi (2006), such discrimination can be highly subjective. Furthermore, Smith et al. (1988) noted an extremely high correlation between news classified as positive and negative. This is due to the fact that media interest drives the volume of news and when coverage increases, both positive and negative news reports rise. A change in the balance between positive and negative news

[9] Both wholesale and consumer prices refer to the whole chicken. The producer price level is not available at a monthly frequency.

[10] While a longer record for retailer prices is available at the monthly frequency, the only available monthly data for wholesale prices are the ones included in the analysis. It was thus not possible to obtain more than 48 observations. Data limitations are a usual problem in analyses focusing on developing countries. While it is better to assess the impacts of food scares on these countries than leaving the question unexplored, data limitations require careful interpretation of our results.

could only be triggered by the disclosure of novel scientific evidence, which rarely happens in the short term. Articles have not been weighted to take into account the size of the article or the location of the article within the newspaper. Although this can be a limitation, this weighting process can be also highly subjective.

Having counted the number of news pieces, the second step is to build the index following Chern and Zuo (1997). The FSII is found to be non sensitive to different values for both the number of lags (n) and peak times (m). Based on these results, we select n=6 which is consistent with the recommendation by Clarke (1976) and requires m to be equal to or less than two. Since the FSII is very similar independently of the chosen value for the peak time, we select m=2. Figure 2 presents the monthly FSII used in our analysis, where it can be observed that before 2006 the FSII is almost zero. Egypt's high self sufficiency rate in chicken meat production may have helped prevent early contamination. The FSII reaches its peak in March 2006, one month after the confirmation of the first Egyptian AI case in February 2006.

A preliminary analysis of the time series data was carried out to assess their time series characteristics. More specifically, standard augmented Dickey and Fuller (1979) tests, Phillips and Perron tests (1988) and Perron tests (1997) were applied to each price series in order to determine whether price series have a unit root. Results are presented in Table 1 and confirm the presence of a unit root in each price series.

Once confirmed that price series are integrated of order one, I(1), we applied the Johansen's (1988) test that supports the existence of a long-run relationship between wholesale and consumer price series (see Table 2). Weak exogeneity tests for long-run parameters were performed in the framework of this method and provided evidence that consumer prices are weakly exogenous. These results are in line with previous literature that has shown that while producer prices tend to adjust to their long-run parity, consumer prices are more sticky or slowly-responsive to price changes occurring in other levels of the marketing chain (Peltzman, 2000; Ben Kaabia and Gil, 2007; Goodwin and Holt, 1999; Serra and Goodwin, 2003; Hassouneh et al., 2010). The equilibrium relationship is thus normalized by the wholesale price. Hansen and Johansen's (1999) test for constancy of the cointegration parameters is also applied and suggests constancy of these parameters throughout the period studied.[11] Saghaian et al. (2008) have also found evidence of cointegration between producer and consumer prices in the Turkish poultry sector.

As already mentioned, in order to apply a STVECM, a linear vector error correction model is first estimated to determine the optimal number of lags in specifying short-run price dynamics. AIC and SBC criteria recommend the use of one lag. Before estimating a non-linear model, a test for linearity was conducted which suggested that linearity is strongly rejected against a STVECM.[12] Initial tests for STVECM specification support the use of an exponential function as the transition function (Teräsvirta, 1994).

After initial specification tests, we proceeded to estimate the parameters of the ESTVECM model by using non-linear least squares (NLS).[13] Results are presented in

[11] Engle and Granger (1987) cointegration test is also performed and suggests that the null hypothesis of no cointegration can be rejected at the 95% confidence level. Results are available from the authors upon request.

[12] Results are available from the authors upon request.

[13] We use $\gamma = 0.5$ as a starting value for γ and values of s_{t-d} close to its mean as a starting value for the threshold parameter. The estimates of the linear model are used as starting values for the rest of the ESTVECM parameters. It is also important to note that all observations are included in the estimation of

Table 3. Relevant parameters in our non-linear ESTVECM are the speed of adjustment to disequilibrium from the long-run parity (vectors $\alpha_i, i = 1, 2$), the speed of transition parameters (γ), the threshold parameter (c) and the short-run parameters λ_i that measure the impact on prices of an increase in the number of news articles.

Our results allow for distinguishing between two different price behaviour regimes (see Figure 2). The first regime, characterized by values of $G(s_{t-d};\gamma,c)$ below 0.5, represents relatively calm periods when the crisis had not yet affected Egypt and news published on the topic only concerned AI affecting other countries. The second regime corresponds to $G(s_{t-d};\gamma,c)$ values above 0.5 and represents price behaviour when Egyptian markets were deeply immersed in the food scare crisis. In this latter regime, Egyptian newspapers widely reported on AI infecting poultry and humans in Egypt.

The speed of adjustment parameters (α) in the second regime representing turbulent markets suggest that, at the 5% significance level, wholesale prices respond to deviations from the long-run equilibrium parity. At the 10% significance level, the response of consumer prices to deviations from equilibrium can also be considered significant. These responses have very different implications for market equilibrium. While wholesale prices move to re-equilibrate the system, consumer prices move in the opposite direction. This suggests that, in times of economic distress, it is likely that retailers make use of their market power to increase their marketing margins. This hypothesis is substantiated by the fact that regulations set forth by the government to respond to the crisis contributed to increase retailers' market power. Among these

the parameters of each regime, but they are assigned different weights (through function G) depending on how close they are to the particular regime being estimated.

regulations is the obligation to slaughter chicken at slaughtering houses instead of retailer shops and using cold chain in chicken transportation. These obligations gave big retail chains the opportunity to gain more market share, benefit from their market power, and increase their marketing margins.

When markets are relatively calm ($G(s_{t-d};\gamma,c)<0.5$), wholesale and consumer prices do not adjust to deviations from the long-run equilibrium.[14] The absence of the response could be explained by both calm markets before the advent of the crisis and the high self-sufficiency rate of chicken meat in Egypt, which was around 100%. This high self-sufficiency rate increased consumer confidence and prevented Egyptian markets from adjusting once Egyptian newspapers reported the presence of AI in other countries.

Compatible with previous research findings focusing on the impacts of food scares in developed economies, we find that the economic impacts of AI in Egypt depend on the magnitude of the food scare as measured by the FSII. As noted, error-correction only takes place in cases when perceptions of the crisis are relevant. Hence, while the characteristics of Egypt, such as a high level of illiteracy and a low income that does not allow most people to buy newspapers, could cast doubt on the capacity of news indices to identify the impacts of food scare crises in this country, our empirical results provide evidence against these arguments. The high correlation between news in different information channels is the most plausible explanation for FSII being able to identify the economic impacts of the AI food scare.

[14] These results are consistent with the concept of threshold error correction explained in Obstfeld and Taylor (1997).

The short-run food scare index parameters (λ_i) indicate that when markets are relatively calm both prices respond negatively to an increase in consumer awareness of the AI crisis. Parameters also indicate that when $G(s_{t-d};\gamma,c) > 0.5$ wholesale prices increase in tandem with an increase in the number of news. The response of consumer prices is also positive though not statistically significant. Results are also consistent within the Egyptian context. The negative response of both consumer and wholesale prices to FSII increases when markets are calm is expected. The positive adjustment of prices when the transition function is $G(s_{t-d};\gamma,c) > 0.5$ is due to the supply shortage caused by the AI-triggered exit of many small and inefficient farms. This supply shortage, together with the recovery of consumption, resulted in higher prices relative to the pre-crisis period.

The threshold parameter (c) is statistically significant at the 95% confidence level and is equal to 27.62. This suggests a symmetric adjustment around a FSII on the order of 28 (see Figure 2). The speed of transition parameter (γ) is statistically significant and equal to 0.64. Since the parameter γ can take any positive value (from zero to infinity), an estimate of 0.64 can be considered as a relatively slow transition speed and thus the transition from $G(s_{t-d};\gamma,c) = 0$ to $G(s_{t-d};\gamma,c) = 1$ occurs very smoothly. This smooth transition contrasts with the impacts of food scares in Europe and North America that often explode on markets rather quickly. Figure 2 presents the evolution of the transition function over time, together with the transition variable and the estimated threshold. The figure illustrates that before 2006 regimes are associated with relatively low values of transition function. The highest transition function values are observed when the AI enters Egypt.

To better understand the potential impact of the AI crisis on wholesale and consumer prices, an Impulse Response Analysis is carried out. Impulse Response Functions (IRFs) provide an effective way to predict the time path of the response of a variable to shocks to the system (Lütkepohl, 2005). More specifically, we derive the IRFs showing the consumer and wholesale price adjustments to an increase in the FSII. The forecast horizon is set to H=8 (the time span that the Egyptian avian market required to respond to the crisis – see Figure 1). Results are presented in Figures 3 to 12. Different magnitudes for the increase in the FSII are considered: 1, 10, 20 and 30. A small increase in the FSII on the order of 1 may occur both in the equilibrium and disequilibrium regimes. Its effects are thus studied within the two price regimes. Two points are selected for initialization: one representing the market before the advent of the crisis (April 2004) and another within the crisis period (April 2006). Bigger FSII increases on the order of 10, 20 and 30 mainly took place during the crisis and their impacts are thus assessed within the disequilibrium regime, taking April 2006 as the initialization point.

Consistently with the parameter estimate discussion presented above, a unit increase in the FSII when the market is calm causes a decline in both wholesale and consumer prices (see Figures 3 and 4, respectively). The decline is of a very small magnitude (below 5% by the end of the adjustment period). Figures 5 and 6 illustrate the results of simulating wholesale and consumer price responses, respectively, to a unit increase in the FSII during turbulent times. The increase generates a marked volatility in prices. This volatility eventually leads to higher wholesale and consumer price levels (the price increase is on the order of 14%). This price increase is consistent with the already mentioned supply shortage caused by the AI crisis, coupled with the recovery of consumption after an initial decline that resulted in a new market equilibrium characterized by higher price levels (see Figure 1). Further, Figures 5 and 6 show that wholesale and consumer prices follow similar patterns as a response to the shock, which reassures the existence of error-correction mechanisms within this price-regime.

Increases of the FSII of a higher magnitude (Figures 7 to 12) generate results similar to the ones discussed above, but of a bigger size. For example, an increase in the FSII of 30 implies an increase in wholesale and consumer prices on the order of 260%. These results accurately reflect what happened during the crisis: from March to April 2006 there was an increase of the FSII on the order of almost 30, which involved an increase in wholesale and consumer prices between 250-260%.

6. Concluding remarks

In this paper we assess the effects of AI on price transmission along the Egyptian poultry marketing chain. We also test whether information on the disease disseminated by mass media is relevant to explain the economic impacts of the crisis. While food scare information indices built upon a count of newspaper articles have been widely used to assess the economic effects of food contamination outbreaks in developed countries, this is the first study that focuses on a developing country.

Following the methodology proposed by Chern and Zuo (1997) we build our FSII upon a monthly count of newspaper articles published in the most widely read Egyptian newspaper. Our empirical model utilizes two series of monthly poultry prices (wholesale and consumer) and one monthly FSII series, which is used as transition variable within a STVECM that allows for nonlinear and smooth price behaviour.

Cointegration tests provide evidence of a long-run equilibrium relationship between poultry prices at different levels of the food marketing chain. The estimated ESTVECM suggests that both wholesale and consumer prices respond to deviations from the long-run equilibrium relationship when the magnitude of the AI crisis is relatively large. Responses of these prices have, however, very different implications for market equilibrium: while wholesale prices respond to correct deviations from the equilibrium, consumer prices respond to increase retailers' marketing margins. This may increase the likelihood of wholesalers abandoning the sector when the magnitude of the crisis is substantially high and probably increases the need for public intervention if this is to be prevented. Conversely, when markets are relatively calm, disequilibriums from the long-run parity do not elicit price responses. These results are expected and are compatible with previous research (Hassouneh et al., 2010).

Results also suggest that food safety information indices representing the degree of food scare in developing countries are useful tools to explain the economic impacts of food crises in these countries. An interesting research finding is that an increase in news when markets are relatively calm has a negative effect on prices, which can help to recover consumption. The effects of an increase in news when markets are already immersed in the crisis results in a strong increase in price levels and volatility. This increase is explained by the supply shortage that AI caused in Egypt coupled with the recovery of consumption after an initial decline. This has important policy implications, since any news issued to recover consumer confidence and demand and thus prevent relevant negative economic effects from a food scare, may be more productive when circulated at the very beginning of the crisis, rather than later on when consumers are already confused and over-saturated by media.

Our analysis can be extended in several ways. One future research line could consist of analyzing the effect of AI on meat demand in Egypt. It would be interesting to test for the relevance of FSII in explaining demand changes. However the availability of the data needed to conduct this type of analysis is not guaranteed. Another interesting research opportunity would be to investigate the effect of food scares in other developing countries by using FSII. This would allow us to test to what extent our results can be generalized.

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http://www.who.int/csr/disease/avian_influenza/country/cases_table_2010_02_17/en/index.html. Accessed 1 March 2010.

Figure 1. Monthly poultry prices.

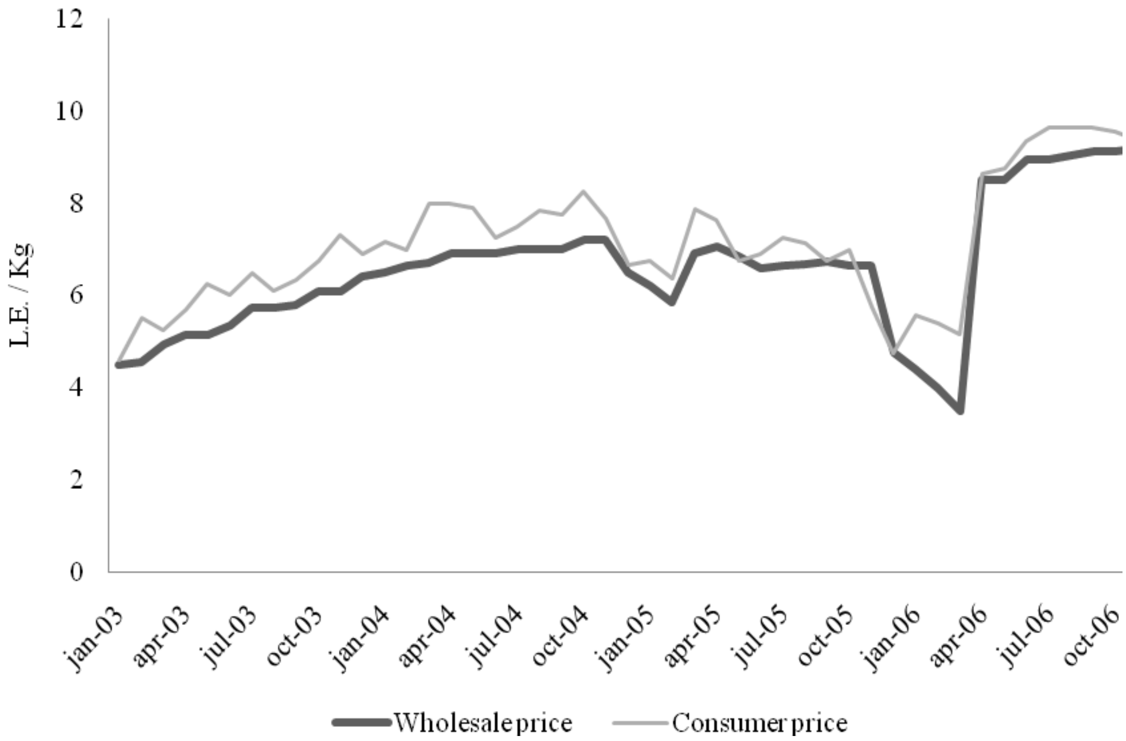
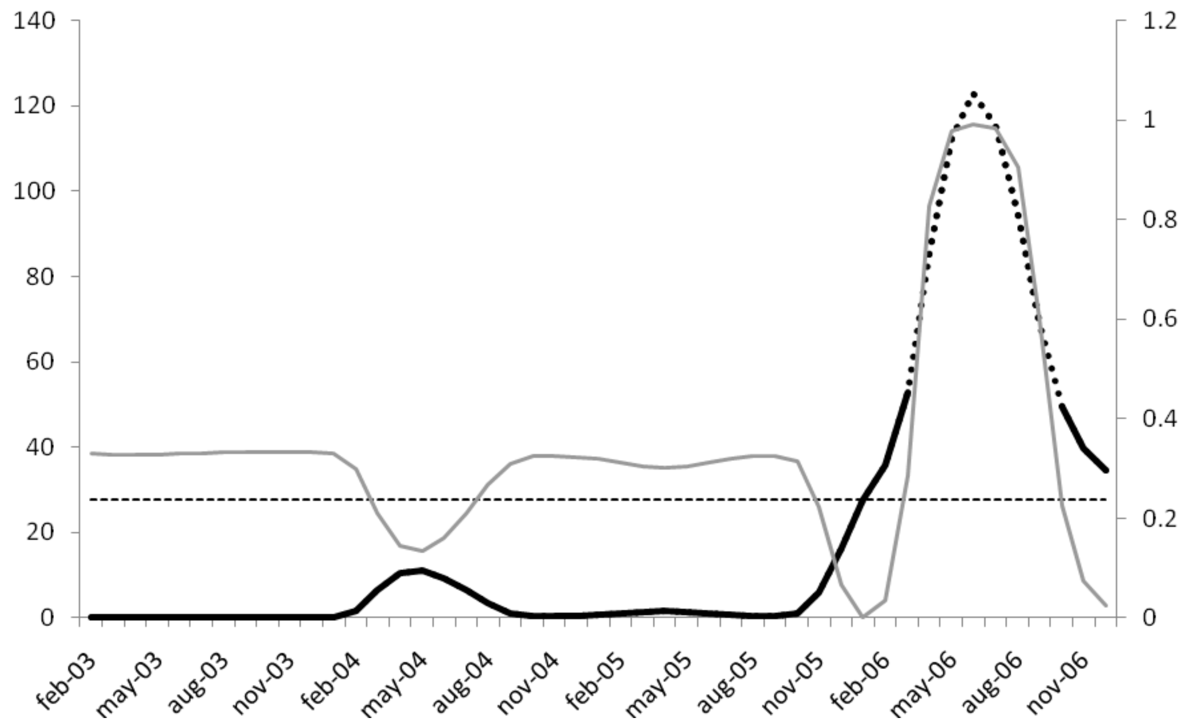


Figure 2. Evolution of the food scare index, threshold variable and the value of the transition function over time.



Notes: The transition function ($G(s_{t-d};\gamma,c)$) is represented by the silver line and is plotted on the right hand side axis. The transition variable (s_{t-d}), presented by the continuous thick black line when $G < 0.5$ and by the dotted thick black line when $G > 0.5$, and the threshold value (thin dotted line) are plotted on the left hand side of the axis.

Figure 3. Wholesale response to a FSII increase of 1. Equilibrium regime.

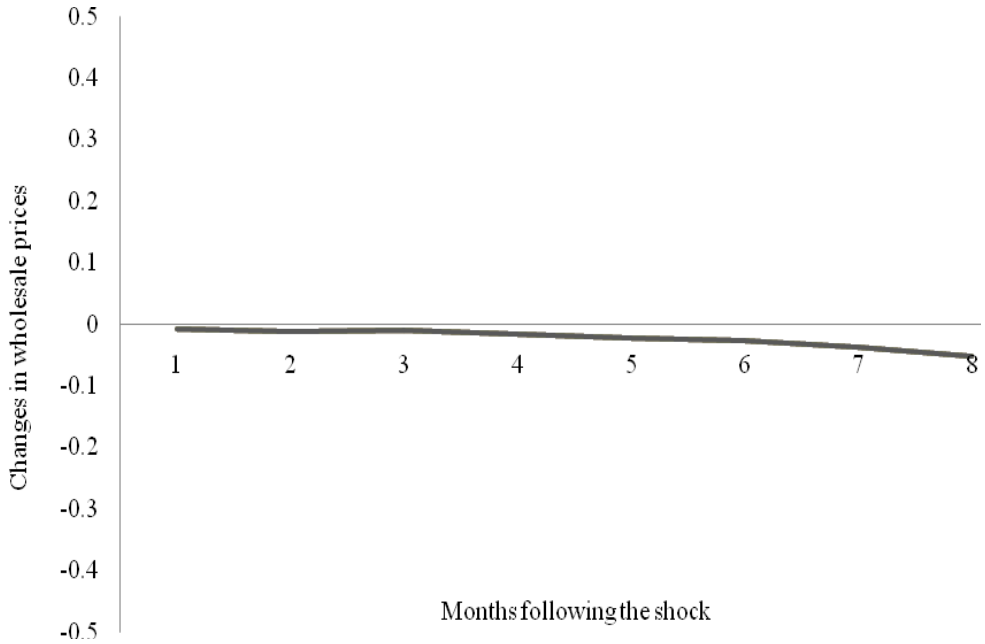


Figure 4. Consumer response to a FSII increase of 1. Equilibrium regime.

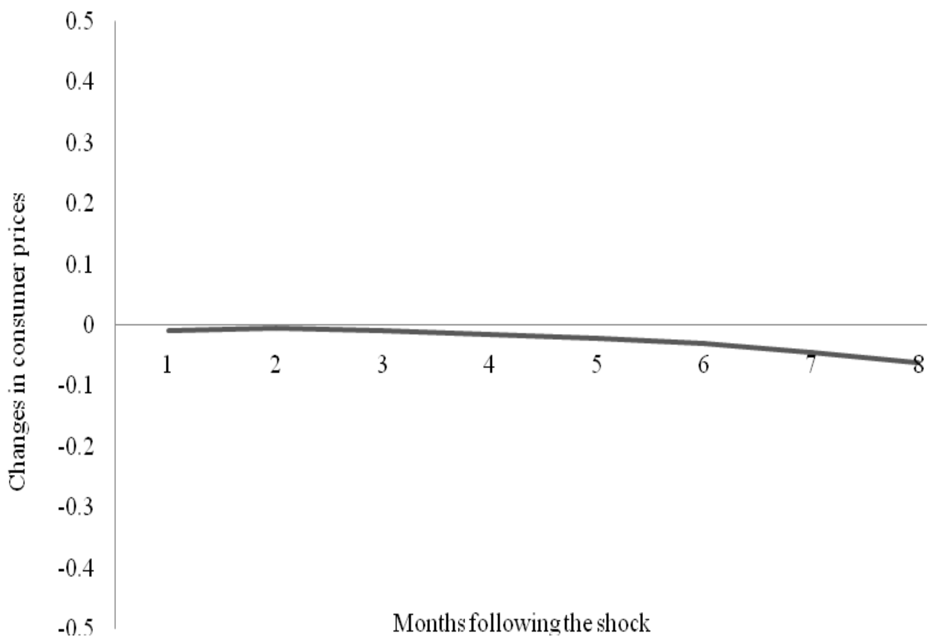


Figure 5. Wholesale response to a FSII increase of 1. Disequilibrium regime.

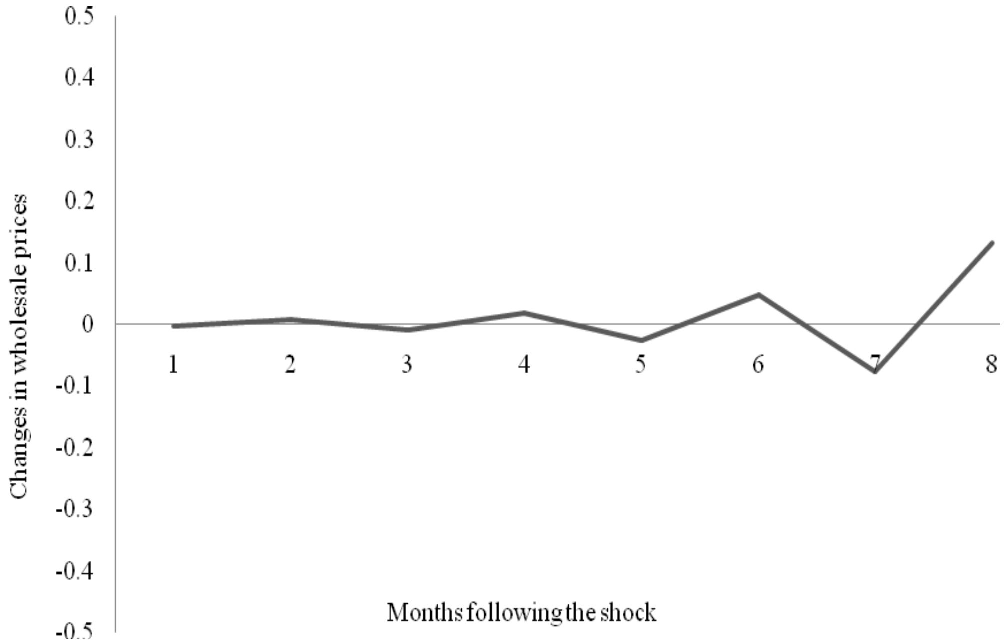


Figure 6. Consumer response to a FSII increase of 1. Disequilibrium regime.

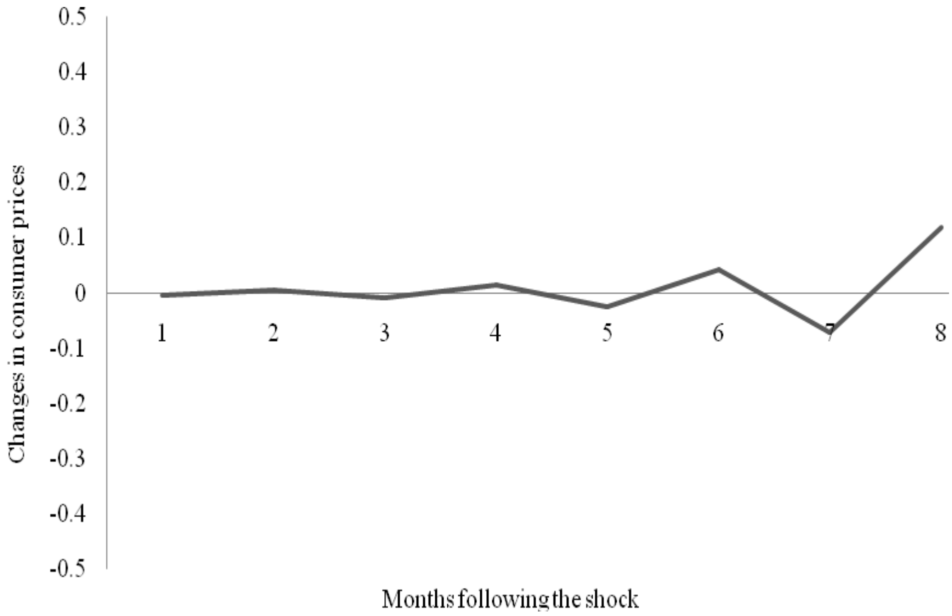


Figure 7. Wholesale response to a FSII increase of 10. Disequilibrium regime.

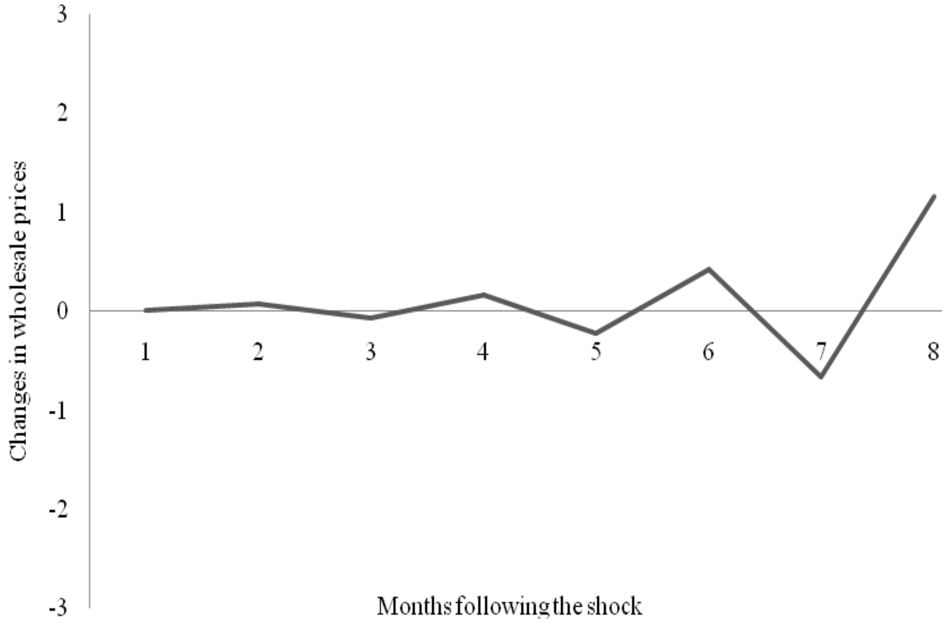


Figure 8. Consumer response to a FSII increase of 10. Disequilibrium regime.

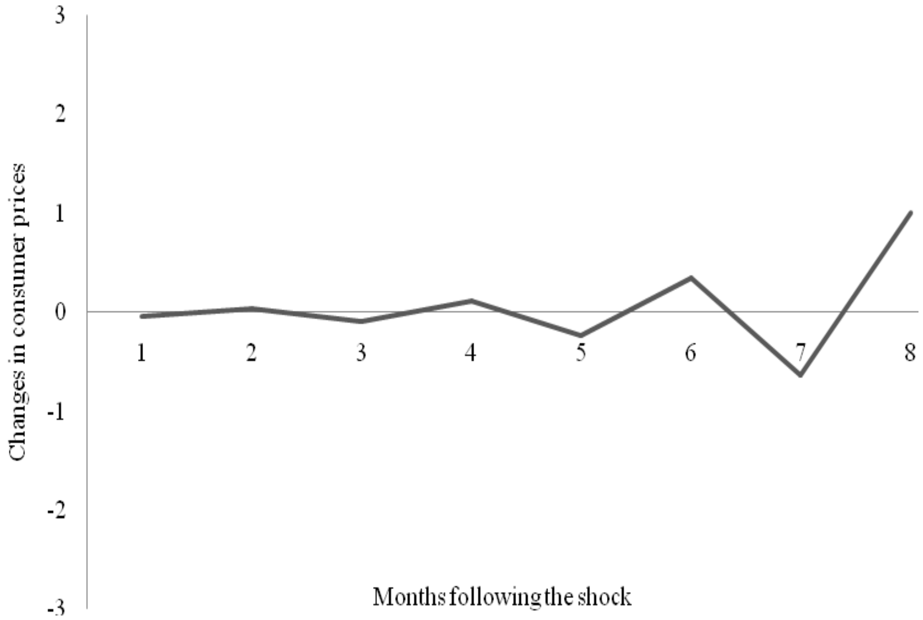


Figure 9. Wholesale response to a FSII increase of 20. Disequilibrium regime.

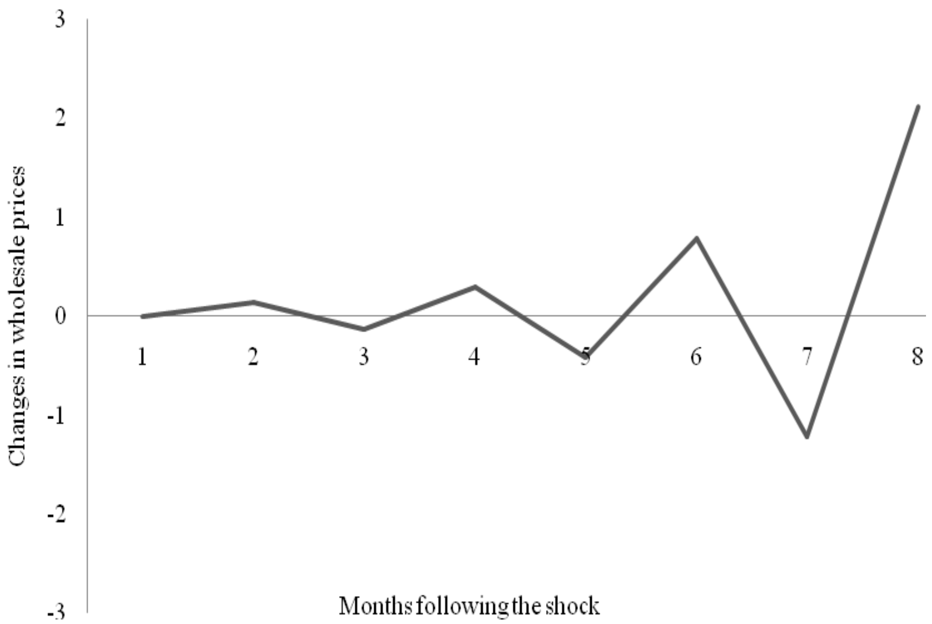


Figure 10. Consumer response to a FSII increase of 20. Disequilibrium regime.

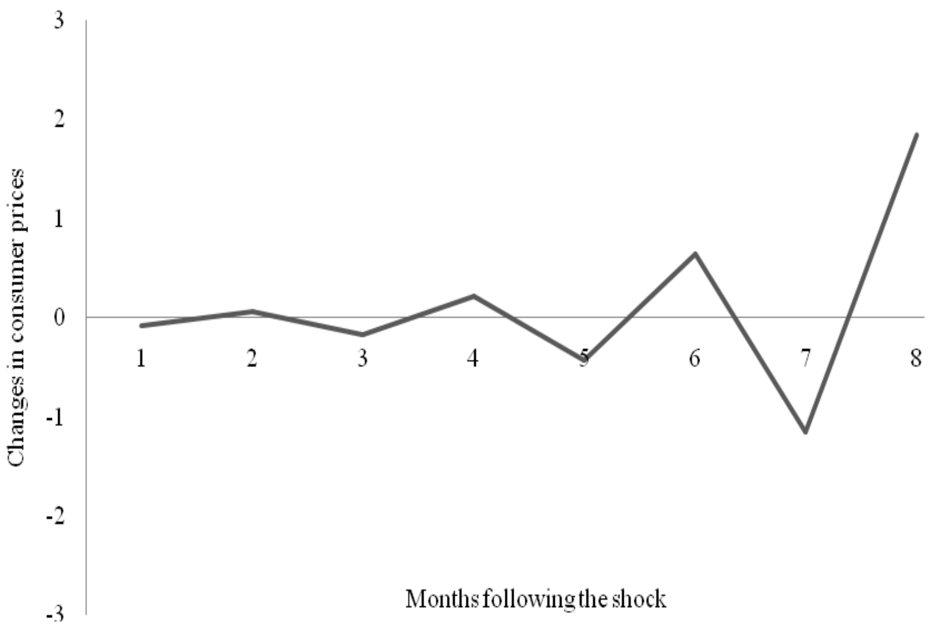


Figure 11. Wholesale response to a FSII increase of 30. Disequilibrium regime.

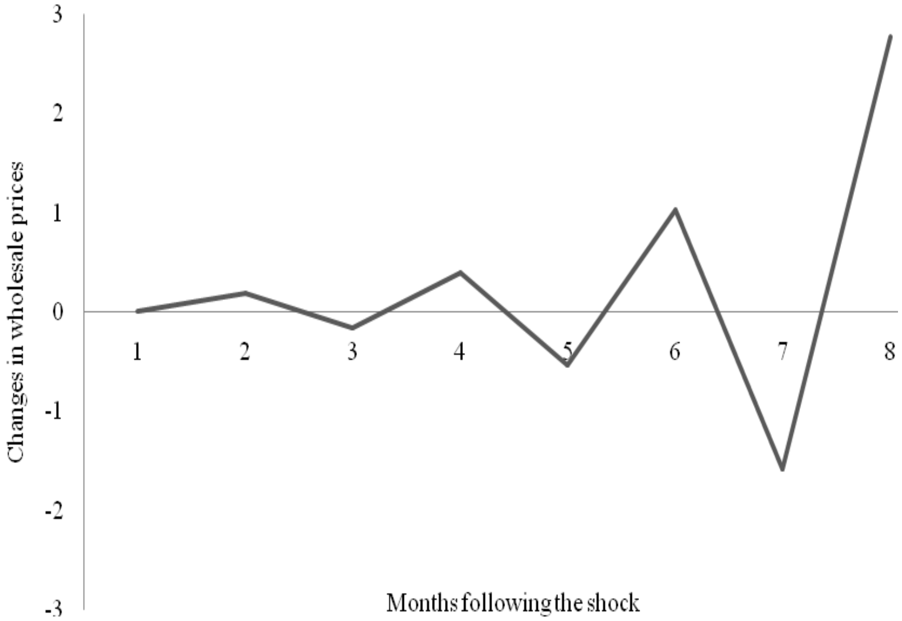


Figure 12. Consumer response to a FSII increase of 30. Disequilibrium regime.

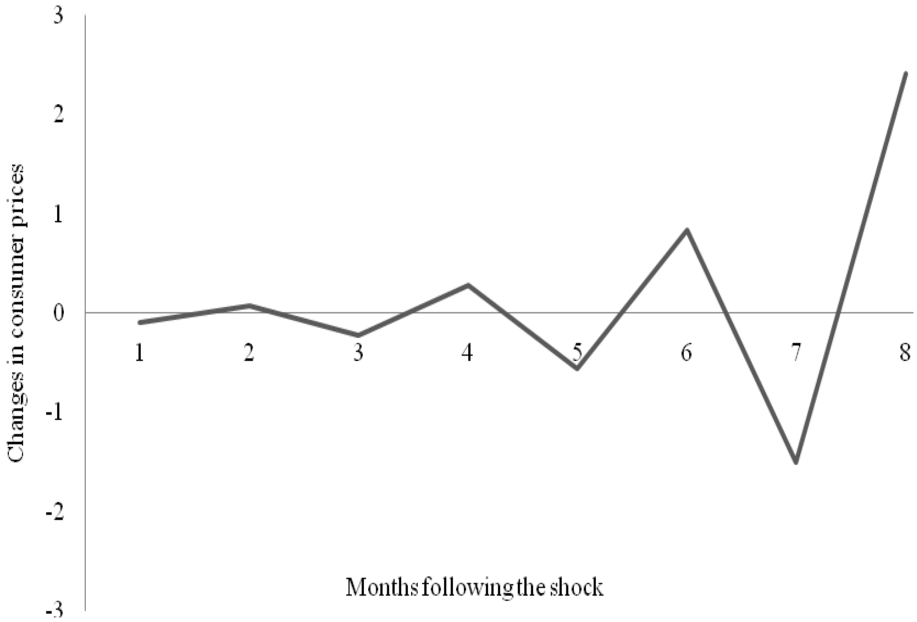


Table 1. Unit root tests.

Price series	Test type	Test statistic (lag)	5% critical value
Wholesale	ADF	-2.687 (0)	-3.447
	PP	-2.690 (0)	-3.447
	Perron	-4.198 (12)	-5.59
Consumer	ADF	-2.578 (0)	-3.447
	PP	-2.580(0)	-3.447
	Perron	-4.523 (3)	-5.59

Notes: The test statistic includes a constant and trend. Critical values for the ADF and PP are obtained from MacKinnon (1991). Critical values for the Perron test are obtained from Perron (1997).[15]

[15] Results of the tests are maintained when using only a constant.

Table 2. Johansen λ_{trace} test for cointegration and cointegration relationship.

Ho	Ha	λ_{trace}	P-value
$r = 0$	$r > 0$	30.841	0.001
$r \leq 1$	$r > 1$	3.973	0.428

Cointegration relationship

(standard errors in parenthesis)

$p_w - 1.136^{**} p_c + 0.348^{**} = 0$

(0.064) (0.126)

Note: r is the cointegration rank.

** denote statistical significance at the 5 % level.

Table 3. ESTVECM parameter estimates.

	Regime G=0 (i=1)		Regime G=1 (i=2)	
Parameters	Parameter estimate	Standard error	Parameter estimate	Standard error
Wholesale price equation				
μ_i	-0.083**	0.025	0.239**	0.078
α_i	0.555	0.417	-2.463**	0.589
Φ_{i,P_w}	0.704**	0.262	1.058*	0.581
Φ_{i,P_c}	1.149**	0.479	-2.210**	1.005
λ_i	-0.011**	0.003	0.011**	0.004
Consumer price equation				
μ_i	-0.060*	0.030	0.181**	0.077
α_i	0.610	0.448	-1.379*	0.766
Φ_{i,P_w}	-0.959**	-0.331	1.241	0.885
Φ_{i,P_c}	0.883	0.557	-2.443	1.483
λ_i	-0.007**	0.003	0.006	0.004
Transition function				
γ	0.645**	0.271		
c	27.622**	1.650		

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

Chapter 4

Non-parametric and parametric modeling of biodiesel, sunflower oil, and crude oil price relationships

1. Introduction

The rapid growth in worldwide biofuel production over the last decade has been mainly motivated by an array of policies targeting different objectives such as the reduction of the energy dependence on fossil fuels, the diversification of energy supply sources, the promotion of economic development in rural areas, or the reduction of greenhouse gas emissions and other sources of environmental degradation (Rajagopal and Zilberman, 2007). The world biofuel market is dominated by biodiesel and ethanol. While ethanol is mainly produced by the US and Brazil, the EU is considered to be the major producer of biodiesel in the world, accounting for about 65% of global output. Within the EU, Spain is the third largest EU biodiesel producer, behind Germany and France (EBB, 2010). Although biofuels may reduce foreign energy dependence, improve rural economies and achieve environmental goals, increased use and production of biofuels poses important challenges.

The major crops used in the biofuels industry are agricultural commodities that can also be used to produce food. In Spain, for example, sunflower oil represents the most relevant raw material used for biodiesel production. The demand for sunflower oil to produce biodiesel directly competes with its use to produce food or animal feed. This raises social and political concerns about the relationship between food and energy prices. It is thus important to determine the extent to which energy (biodiesel) markets have the potential to increase food commodity (e.g., sunflower oil) prices.

Our paper sheds light on this issue. While the previous literature has focused on studying price linkages and price transmission patterns between energy and feedstock markets in the US and Brazil (see, for example, Balcombe and Rapsomanikis, 2008;

Zhang et al., 2009; Serra et al., 2010 and 2011), European markets have not yet received any direct research attention. The objective of this research is to fill this gap in the literature by assessing the relationship between biodiesel, sunflower and crude oil prices in Spain, one of the most relevant biodiesel producers within Europe.

To achieve the aforementioned objective, two different time series models are considered. First, we estimate a parametric Vector Error Correction (VEC) model that studies the short-run and long-run dynamics of price transmission among the markets considered. This parametric approach requires assumptions about the nature of price behavior that may be too restrictive and thereby may generate misleading results. More specifically, VEC models are based upon the assumption that price linkages are of a linear nature. In order to study to what extent the linearity hypothesis and the results obtained from the parametric VECM are reliable, a non-parametric model is also fit to the data. Unlike parametric methods, non-parametric techniques such as Multivariate Local Polynomial Regression (MLPR) methods, are data driven and do not require any specific assumption about the nature of price transmission. Thus, they are robust to misspecification issues. To our knowledge, no previous study has utilized non-parametric modeling to assess the relationships between energy and feedstock prices. This application represents another important contribution of our work to the existing literature.

The rest of the paper is organised as follows. Section 2 presents an overview of the EU and the Spanish biodiesel industry. After offering a brief review of the literature, in section 4 we describe both the parametric and non-parametric methods used to assess the nature of price transmission. The results are presented in section 5. Finally, the paper ends with the concluding remarks section.

2. The EU and the Spanish biodiesel industry

Recent increases in crude oil prices as well as the fact that oil reserves are limited have been the main drivers for industrialized countries to promote alternative energy sources such as biofuels to diversify energy supplies, and thereby reduce dependency on fossil fuels. Biofuels are transportation fuels that are mainly produced from agricultural inputs. Biodiesel and ethanol are the most common biofuels utilized in the transportation sector worldwide.

In the EU, biofuels have experienced dynamic development over the recent years. Biofuel use for transportation in the EU grew by 30.3% between 2007 and 2008. In 2009, consumption continued to increase, but at a slower rate than in 2008 (18.7%). As a result, the share of biofuels in total EU consumption of road transportation fuels increased from 3.3% in 2008 to 4% in 2009. More specifically, in 2009 the EU consumption of biofuels reached almost 12.1 million tons of oil equivalents (toe), of which 9.6 million tons were biodiesel and 2.3 million tons were ethanol. The predominance of biodiesel is due to the relevance of diesel consumption in the EU compared to gasoline (EurObserv'ER, 2010).[1]

With a consumption of almost 1.05 million toe in 2009, of which 894 thousand toe were biodiesel and 152 thousand toe were ethanol, Spain was the fourth largest EU consumer of biofuels, behind Germany, France and Italy. Biofuels consumption growth has mainly involved increased use of biodiesel, which role by almost 72% in 2009 compared to 2008.

[1] Unless otherwise indicated, the information presented in this section was obtained from EurObserv'ER (2010).

The implementation of the EU biofuels policy, which aimed at replacing, by the end of 2010, 5.75% of the non-renewable energy used in transportation by biofuels (Directive 2003/30 EC), as well as the recent increases in crude oil prices, have positively affected the development of the EU biofuels industry. This has resulted in the increase of both EU biofuels production and productive capacity. The expansion has been especially evident in the biodiesel industry.

The EU biodiesel industry plays a major role both in the EU and at the international level. According to the European Biodiesel Board (EBB, 2009), EU biodiesel production capacity increased from 16 to 21 million tons between 2008 and 2009 (a 31% increase). In spite of the progress made, the EU industry remains uncompetitive in relation to the American and Brazilian industries, which have much lower production costs. The low price of the biodiesel imported from the US causes significant damage to the EU biodiesel industry, particularly in terms of profitability and return on investments. To address increasing international competition, duties on imports of US biodiesel were applied in 2009.

Nowadays, the EU is considered to be the biggest producer of biodiesel in the world, representing about 65% of worldwide output (EBB, 2010). Biodiesel production rose by 16.6% (9 million tons) between 2008 and 2009, against 35.7% (7.7 million tons) between 2007 and 2008 and 54% (5.7 million tons) between 2006 and 2007, leaving almost half of the production capacity unused.[2] In 2009, biodiesel represented almost 75% of the biofuels produced in the EU. Despite biodiesel production being below capacity, the EU's renewable energy policy that aims to ensure that fuels sold in the EU

[2] This unused production capacity can be explained by the considerable investments in biodiesel plants planned already before 2007, as a response to the ambitious goals for renewable energy use within the EU countries (EBB, 2010).

contain a minimum 20% of renewable energy by 2020, as well as the duties against US biodiesel imports, are expected to positively affect the EU biofuels industry.

In Spain, according to the EBB (2010), biodiesel production increased from 207 thousand tons in 2008 to 859 thousand tons in 2009, making Spain the third largest biodiesel producer in the EU after Germany and France. Currently, there are 46 biodiesel production plants in Spain with a total productive capacity of over 4.2 million tons, leaving almost four-fifths of the productive capacity unused (APPA, 2010a). This excess capacity can be mainly explained by the expectations created by the Directive 2003/30 EC (APPA, 2010b). In contrast to the EU as a whole, where the majority of the biodiesel is produced from rapeseed oil, sunflower oil constitutes the major input of the biodiesel industry in Spain.

In 2009, Spanish oilseed production registered an increase of 5%, reaching a level of 886 thousand tons, of which 851 thousand tons were sunflower oil, 32 thousand tons rapeseed oil and 3 thousand tons soybean oil, respectively (Agrocope, 2009). The area used to produce oilseeds in 2009 increased by 15%, representing 871 thousand hectares, of which, 852 thousand hectares were used for sunflower production (Alimentación en España, 2009).

The recent rapid growth in biodiesel production registered both in the EU and in Spain have raised concerns about the impact that biodiesel may have on agricultural commodity prices. In this paper, we focus on assessing the linkages between food and energy prices within the Spanish biodiesel industry.

3. Literature review

Even though a large number of studies and reports on biofuels have been made available recently, the literature that empirically assesses the dynamic relationships between energy and commodity prices is relatively poor. A few notable studies on this topic are reviewed below. Balcombe and Rapsomanikis (2008) examine the nature of the relationships between crude oil, ethanol and sugar prices in Brazilian markets. A Bayesian approach is used to test for non-linear price-adjustments. They find that oil prices are the main drivers of both sugar and ethanol prices. Specifically, they show that a causal hierarchy runs from oil to sugar and to ethanol, and that non-linearities characterize price adjustment processes of sugar and ethanol prices to the oil price.

Zhang et al. (2009) investigate the impact that ethanol markets have on price levels and volatilities of agricultural commodities in the US. In doing so, they use cointegration, vector error correction and Multivariate Generalized Autoregressive Conditional Heteroskedascity (MGARCH) models. Their results indicate that, although short-run linkages exist, no long-run relationships between fuel and food prices exist.

Serra et al. (2010) utilize a Smooth Transition Vector Error Correction Model (STVECM) in order to study the relationships among oil, ethanol, gasoline and corn prices within the US market. They find two equilibrium relationships among the four prices studied. Adjustment towards long-run equilibrium is found to occur in a non-linear fashion. Serra et al. (2010) further find that an increase in energy prices increases corn prices, meaning that the ethanol market is mainly responsible for the strong link between food and energy prices.

The analysis by Serra et al. (2011) studies how price volatility in the Brazilian ethanol industry changes over time and across markets. Seo's (2007) maximum likelihood approach, which allows estimating a VECM and a MGARCH model jointly, is used for such purpose. Their results suggest a strong link between food and energy markets, both in terms of price levels and volatilities. The use of flexible semiparametric GARCH techniques allows Serra (2011) to refine the conclusions derived in the previous analysis.

The existing literature has mainly assessed the dynamic linkages between food and energy markets for the major ethanol markets in the world, (i.e., Brazil and the US). Emergent biofuel markets in Europe have, however, been ignored in existing research. Our paper fills this gap in the existing literature by focusing on Spain, a relevant biofuels market within Europe.

Unlike the existing literature, which reveals that analyses of the dynamic relationships among energy and agricultural commodity prices have been typically based on parametric models, we are interested in studying the relationships among biodiesel, sunflower and crude oil prices in Spain by using multivariate, local, polynomial regression (non-parametric) techniques. Non-parametric techniques are data-driven methods that do not require any assumptions about the functional form characterizing price behavior and are thus robust to misspecification issues (Fan and Gijbels, 1996; Li and Racine, 2007). The results of non-parametric techniques are compared with those derived from a parametric VEC model.

4. Methodology

To study the relationships between biodiesel, sunflower and crude oil prices in Spain, we use both a parametric VECM and a non-parametric MLPR model and compare the results obtained from each. An interesting issue is the complementarity between non-parametric and parametric techniques. Specifically, MLPR and VEC methods can be applied in similar situations, but they serve different analytical purposes. While the parametric approach focuses on estimating and making inferences that are based on parameters of the model, the MLPR method concentrates on exploring data non-parametrically, which can help visualize data effects and can thus be considered as a powerful tool for data exploration and features that might otherwise be missed. It is thus interesting to apply both parametric and non-parametric estimation techniques.

4.1 Vector Error Correction model

Myers (1994) explains that price series have different common characteristics that are important for sound statistical analysis. Two of these characteristics are especially relevant to our analysis. First, individual commodity price series generally contain stochastic trends and therefore are non-stationary. Second, commodity prices may tend to move together over time. In other words, though individual price series may be non-stationary, price series of interrelated markets are likely to contain the same stochastic trends. Hence, the co-movements of these variables may be stationary. Co-movement among non-stationary prices is known in the econometrics literature through the concept of cointegration. VEC models allow assessing both short-run price dynamics and the adjustment of individual prices to deviations from the long-run cointegration relationship.

Assume that equation (1) represents the cointegration relationship between the prices studied:

$$P_{B,t} - \beta P_{S,t} - \beta P_{C,t} = v_t \tag{1}$$

where $P_{B,t}$, $P_{S,t}$, $P_{C,t}$ are the prices of biodiesel, sunflower oil and crude oil at time *t*, respectively and v_t represents the deviation from the equilibrium relationship, i.e., the error correction term. If the series are found to be cointegrated, and following Engle and Granger (1987), a VECM can be expressed as follows:

$$\begin{aligned}
\Delta P_{B,t} &= \alpha_1 + \lambda_B v_{t-1} + \sum_{i=1}^{n} \alpha_{11}(i)\Delta P_{B,t-i} + \sum_{i=1}^{n} \alpha_{12}(i)\Delta P_{S,t-i} + \sum_{i=1}^{n} \alpha_{13}(i)\Delta P_{C,t-i} + \varepsilon_{P_{B,t}} \\
\Delta P_{S,t} &= \alpha_2 + \lambda_S v_{t-1} + \sum_{i=1}^{n} \alpha_{21}(i)\Delta P_{B,t-i} + \sum_{i=1}^{n} \alpha_{22}(i)\Delta P_{S,t-i} + \sum_{i=1}^{n} \alpha_{23}(i)\Delta P_{C,t-i} + \varepsilon_{P_{S,t}} \\
\Delta P_{C,t} &= \alpha_3 + \lambda_C v_{t-1} + \sum_{i=1}^{n} \alpha_{31}(i)\Delta P_{B,t-i} + \sum_{i=1}^{n} \alpha_{32}(i)\Delta P_{S,t-i} + \sum_{i=1}^{n} \alpha_{33}(i)\Delta P_{C,t-i} + \varepsilon_{P_{C,t}}
\end{aligned} \tag{2}$$

where Δ is a first difference operator; $\varepsilon_{P_{B,t}}$, $\varepsilon_{P_{S,t}}$ and $\varepsilon_{P_{C,t}}$ are white noise disturbances; v_{t-1} is the lagged error correction term, the α terms are all short-run dynamics parameters; and λ_B , λ_S and λ_C, known as the speed of adjustment parameters, measure the rate at which prices adjust to disequilibria from the long-run equilibrium relationship. It is important to note that all the variables in the VECM are stationary by virtue of cointegration relationships and first-differencing transformations.

Parameters λ_B , λ_S and λ_C offer particularly valuable information and should not be simultaneously equal to zero. In particular, at least one of these speed of adjustment terms must be nonzero for a long-run equilibrium relationship between prices to exist. If λ_B , λ_S and λ_C are all equal to zero, the long-run equilibrium relationship does not exist. Nevertheless one may find that only one of the speed of adjustment parameters is significantly different from zero, indicating that there is only one price that does all the adjustment towards the long-run equilibrium. The VEC model can be estimated by the seemingly unrelated regressions technique.

Before estimating the VEC model, standard unit root and cointegration tests were conducted in order to determine whether price series are stationary and whether they are cointegrated, respectively. In particular, standard augmented Dickey and Fuller (1979) tests and Perron tests (1997) were applied to each price series. The Johansen (1988) test for cointegration was then used to evaluate long-run price linkages.

4.2 Multivariate local polynomial fitting

Though parametric models are the workhorse of applied data analysis, they require specifying the exact functional form of the model prior to estimation. A parametric model that is not accurately specified may lead to misleading results. Non-parametric regression models do not impose any restriction on the functional form and thus allow one to explore the data in a more flexible way. Local polynomial techniques are often used to estimate regression functions in a non-parametric fashion, where data are segmented into small overlapping sections (Fan and Gijbels, 1996).

To formally study the relationship between biodiesel, sunflower and crude oil prices in Spain, we apply a multivariate local polynomial regression to estimate a non-parametric version of the VEC model. Three non-parametric error correction equations are estimated, one for each equation in the VEC model. Consider a set of observations $(Y_t, \mathbf{X}_{t-1})$ for $t = 1, ..., n$ from a population $(Y, \mathbf{X}_{-1})$, where $Y_t = \Delta P_{it}$, $Y_t \in \mathbb{R}$, represents price i in first differences, i= B, S, C is an index representing biodiesel, sunflower oil and crude oil respectively, and $\mathbf{X}_{t-1} = (\Delta P_{Bt-1}, \Delta P_{St-1}, \Delta P_{Ct-1}, v_{t-1})$, $\mathbf{X}_{t-1} \in \mathbb{R}^d$ is a vector containing lagged price differences of biodiesel, sunflower and crude oil prices, and the lagged error correction term, being $d = 4$. Of interest is to estimate the multivariate non-parametric regression problem $m(\mathbf{x}_k) = E(Y_t \mid \mathbf{X}_{t-1} = \mathbf{x}_k)$.

The main idea behind local fitting is to estimate the function m at point $\mathbf{x}_k$, i.e. $\hat{m}(\mathbf{x}_k)$, using the observations that are relatively close to $\mathbf{x}_k$. To estimate the entire function $\hat{m}(\mathbf{X}_{t-1})$, the process is repeated for a number of grid values of $\mathbf{X}_{t-1}$ (Serra et al., 2006). Since the function $m(\mathbf{x}_k)$ is not specified, a Taylor series expansion is used to approximate it by a simple polynomial model:

$$m(\mathbf{x}) \approx \sum_{j=0}^{p} \boldsymbol{\beta}_j^{'} (\mathbf{x} - \mathbf{x}_k)^j \tag{3}$$

where the local parameter vector $\boldsymbol{\beta}_j = m^{(j)}(\mathbf{x}_k) / j!$ depends on $\mathbf{x}_k$. The $m^{(j)}$ term is the jth derivative of function m. Previous research argues that odd order polynomial fits are preferable to even order polynomial fits, (see, Fan and Gijbels, 1996 for a discussion on selecting the polynomial order).[3] Moreover, several authors recommend choosing a

[3] Fan and Gijbels (1996) show that odd order fit provides a significant bias correction especially in the boundary areas and in highly clustered reigns.

polynomial of order $p = 1$, which leads to the Multivariate Local Linear Regression Estimator (MLLRE), an estimator that has been shown to provide adequate smoothed points and computational ease (Cleveland, 1979; Heij et al., 2004; Wu and Zhang 2006).

The observations with most information about $m(\mathbf{x}_k)$ should be those at locations closest to $\mathbf{x}_k$ compared to more remote points. Weighted least squares is used to give more weight to neighboring observations than to more distant ones. Weights are assigned through a kernel functions as follows:

$$\sum_{t=1}^{n}(Y_t - \beta_0 - \boldsymbol{\beta}_1^{'}(\mathbf{X}_{t-1} - \mathbf{x}_k))^2 K_{\mathbf{h}}\left(\mathbf{X}_{t-1} - \mathbf{x}_k\right) \tag{4}$$

where $K_{\mathbf{h}}\left(\mathbf{X}_{t-1} - \mathbf{x}_k\right) = \prod_{j=1}^{d} K\left(\frac{X_{j,t-1} - x_{j,k}}{h_j}\right) h_j^{-1}$ is a multivariate multiplicative kernel function assigning weights to each datum point and $K\left(\frac{X_{j,t-1} - x_{j,k}}{h_j}\right)$ is a univariate kernel function. The bandwidth h_j controls for the size of the local neighborhood (Fan and Gijbels, 1996) and is equal to $h_j = h_{base} s_x n,^{-1/5}$ where s_x is the standard deviation of the covariate and n is the number of observations (Serra and Goodwin, 2009).[4] The local linear estimate of $m(\mathbf{x}_k)$ is $\hat{\beta}_0$, while the gradient vector $m'(\mathbf{x}_k)$ is $\hat{\boldsymbol{\beta}}_1$.

Previous literature argues that selecting an appropriate bandwidth is a key issue of multinomial local polynomial fitting. Specifically, selecting a large bandwidth may

[4] This is known as the "rule of thumb" bandwidth estimator which was originally suggested by Silverman (1986).

lead to an important modeling bias, while selecting a small bandwidth may result in noisy estimates. In order to find a balance in the bias-variance trade-off, we select an optimum constant base bandwidth h_{base} using the least squares cross-validation method. This commonly used method (Fan and Gijbels, 1996; Li and Racine, 2007) determines a smoothing bandwidth matrix that leads to the minimization of the squared prediction error: $\sum_{t=1}^{n}(Y_t - \hat{Y}_t)^2$. In our analysis and following Kumbhakar et al. (2007), the predicted values for Y_t are obtained using the "leave one out" version of the local linear estimator.[5]

Another issue in multivariate local polynomial fitting is the choice of the kernel function. Fan and Gijbels (1996) argue that the choice of the kernel function is less crucial since the modeling bias is primarily controlled by the bandwidth. However, they recommend using the Epanechnikov kernel which has been shown to be optimal under general conditions.

5. Results

Our empirical analysis utilizes weekly prices for refined sunflower oil, biodiesel and crude oil observed from November 7, 2006 to October 5, 2010, giving a total of 205 observations. Sunflower, biodiesel and crude oil prices are expressed in euros per 100 kg, euros per litter and dollars per barrel, respectively. Data on refined sunflower oil prices were taken from the Spanish Ministry of the Environment and Rural and Marine

[5] This method has been used by Kumbhakar et al. (2007) and Serra and Goodwin (2009) in a multivariate framework.

Affairs (2010), biodiesel prices were obtained from the Spanish Ministry of Industry, Tourism and Trade (2010) and crude oil prices from the US Energy Information Administration (2010) data set. Crude oil prices were converted from dollars to euros per liter using the European Central Bank (ECB, 2010) exchange rates. Sunflower oil prices were also converted into euros per kg. Price series used in our analysis are presented in Figure 1.

Logarithmic transformations of the price series were used in the empirical analysis. A preliminary analysis of the time series data was carried out to assess their time series characteristics. In particular, standard augmented Dickey and Fuller (1979) and Perron tests (1997) were applied to each price series in order to determine if they have unit roots. Results confirm the presence of a unit root in all logarithmic price series.[6]

Johansen's (1988) method was then applied in order to test for log-run linkages among the prices. Results suggest the existence of a single cointegration relationship between sunflower oil and the energy price series (see Table 1). As will be seen below in the estimates of the VEC model, sunflower and crude oil prices are weakly exogenous for long-run parameters, which implies that they do not react to deviations from the long-run equilibrium. Only the biodiesel price responds to deviations from this parity. Hence, the cointegration relationship should be interpreted as representing the relationship that biodiesel prices should maintain with sunflower and crude oil prices for the Spanish biodiesel industry to be in equilibrium. Long-run relationships between crude oil, ethanol and feedstock prices in the US and Brazil have been identified by

[6] Results are available from the authors upon request.

Balcombe and Rapsomanikis (2008), Serra et al. (2010) and Serra et al. (2011), respectively.

The cointegration relationship suggests positive correlation in the long-run between biodiesel and sunflower and crude oil prices.[7] More specifically, the cointegration relationship suggests that an increase in crude oil (sunflower oil) prices on the order of 10% will be followed by an increase in biodiesel prices on the order of 4% (0.8%). The positive relationship between biodiesel and sunflower oil prices is expected, given that feedstock costs represent a considerable part of biodiesel production costs. Feedstock costs are especially relevant within the Spanish biodiesel industry, that has higher production costs than other more competitive industries such as the US industry. The imposition by the EU countries of import duties on US biodiesel imports in 2009 may contribute to perpetuate the lack of competitiveness of the Spanish market. The long-run positive link between biodiesel and crude oil prices is not surprising either and is due to the fact that biodiesel is not usually used in pure form, but rather is blended with petroleum diesel that comes from refined crude oil.

Results derived from the VECM estimation are presented in Table 2.[8] The coefficients showing price adjustments to the long-run parity suggest that, while biodiesel prices adjust to correct disequilibrium condition, sunflower and crude oil prices do not adjust. The biodiesel price is thus the only variable that responds to

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[7] Hansen and Johansen's (1999) test for constancy of the cointegration parameters is applied and suggests constancy of these parameters throughout the period studied. Furthermore, the fluctuation test of the eigenvalues is carried out and provides evidence of constancy of the eigenvalues at the 5% significance level.

[8] The Akaike Information Criterion (AIC) as well as Schwartz Bayesian Criterion (SBC) are used to select the optimal number of lags in the VECM. Both criteria recommend using one lag.

deviations from the biodiesel, sunflower oil and crude oil prices long-run equilibrium and moves to re-equilibrate the price system (-6.7%). These results are not surprising and are compatible with previous research (Balcombe and Rapsomanikis, 2008; Serra et al., 2011). Further, parameter estimates show that crude oil prices have the capacity to influence biodiesel prices not only through the long-run price dynamics, but also through the short run price links. While energy prices are not found to influence sunflower prices through the long-run equilibrium relationship, they have the capacity to increase sunflower oil prices through the biofuel market by means of short-term price dynamics.

The VECM imposes a linear adjustment of the variables being studied. However, the existing literature on price transmission within biofuel markets has generally allowed for and confirmed the existence of nonlinear price adjustments. Unlike parametric methods, non-parametric regression models are data driven and thus do not make any assumption about the functional form characterizing price links. Multivariate local linear regression is then applied in order to determine whether price dynamics are linear or whether they would be best modelled using a different functional form. The most relevant results are graphed in Figure 2.[9] The figure shows the variation of the local estimates of the parameters representing the adjustment of biodiesel,

[9] The optimal base bandwidths h_{base} are searched using a grid of values ranging between 12 and 60 for following the least squares cross-validation method. Optimal smoothing parameters that minimize the squared error of biodiesel, sunflower and crude oil equations are 23, 46 and 43, respectively. It is important to note that the Nadaraya-Watson non-parametric regression estimator (Nadaraya, 1964; Watson, 1964) was also applied as an alternative technique for bandwidth selection and the results obtained were very similar.

sunflower and crude oil prices to deviations from the long-run parity.[10] Non-parametric results suggest a relatively small variation in parameter estimates, thus supporting the existence of linear price adjustments as modeled by the parametric model. This conclusion is further supported by the fitting of a smooth transition vector error correction model (STVECM) to the data (Luukkonen et al., 1988; Teräsvirta, 1994) that showed that the speed of transition parameter is not statistically different from zero.[11]

We now focus on the parameter showing the adjustment of the biodiesel price in the long-run equilibrium relationship, as the adjustment of the sunflower and crude oil prices was not found to be statistically significant. Non-parametric results indicate that biodiesel price responses to long-run disequilibrium conditions, can range from 6.1% to 8.1%. Figure 3 shows that the magnitude of the adjustment depends on the magnitude and sign of the disequilibrium shock. In general, when the error correction term is positive, the biodiesel price rate of adjustment to long-run equilibrium is on the order of almost 6%. The response is quicker, almost 8%, when the error correction parameter values are lower than zero.

These different responses in biodiesel prices are found to have very important implications for market equilibrium. Positive values of the error correction term suggest that biodiesel is too expensive and thus that its price has to decline for the market equilibrium to be maintained. Conversely, negative error correction values imply that the biofuel is too cheap and its price should increase. Hence, shocks that increase biodiesel prices will trigger quicker responses than shocks that reduce biodiesel prices.

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[10] Results of localized short-run parameters are available from the authors upon request.

[11] Results are available upon request.

Given the reliability of the parametric VECM model as shown by the small variation of localized parameter estimates, and to better understand the dynamic relationships between biodiesel, sunflower and crude oil prices, an impulse response analysis is conducted. Impulse Response Functions (IRFs) illustrate the evolution over time of the response of one variable to a shock in another variable in the system (see Lütkepohl, 2005). Since the crude oil price is exogenous with respect to the error correction term as well as the short-run dynamic parameters, only the IRFs showing the biodiesel and sunflower oil price adjustments to system shocks are presented.

Figures 4 and 5 illustrate the reaction of biodiesel price responses to a positive one standard deviation shock to the crude and sunflower oil prices, respectively. Figure 4 shows that an increase in crude oil prices generates a response in the biodiesel price in the same direction. The response increases during the first two weeks following the shock and decreases thereafter, disappearing after about 8 weeks. A positive sunflower oil price shock is also seen to induce an increase in the biodiesel price. This increase gains strength during the first 3 weeks after the shock and shrinks thereafter, disappearing after about 12 weeks (see Figure 5). Hence, biodiesel producers appear to pass on an increase in feedstock costs more slowly than an increase in the price of crude oil.

Figures 6 and 7 show the responses of sunflower oil to a positive standard deviation shock to the biodiesel and crude oil prices, respectively. An increase in biodiesel price is seen to induce an increase in the sunflower oil price that is especially strong during the first two weeks. The sunflower oil price response tapers off after almost 12 weeks. The response of sunflower oil prices to an increase in crude oil price is also shown to be positive and to disappear after almost 12 weeks (see Figure7).

6. Concluding remarks

Recent increases in biofuel production that have been undertaken to reduce dependence on crude oil, diversify energy supplies, support rural economies and reduce greenhouse gas emissions, have generated social and political concerns with regard to the linkages between energy and food price levels. Currently, food inputs represent the main feedstocks used in biofuel production. In this paper we shed light on this issue by assessing price linkages and price transmission patterns between biodiesel, sunflower and crude oil prices in Spain.

To achieve the aforementioned objective, a parametric vector error correction model and an alternative multivariate local polynomial version are estimated and compared. To the best of our knowledge, no previous study has utilized non-parametric modeling to assess the relationships between energy and feedstock prices. Weekly world crude oil prices, and Spanish biodiesel and sunflower oil prices observed from November 2006 to October 2010 are used in the empirical analysis.

Cointegration tests provide evidence of a single long-run equilibrium relationship among biodiesel, sunflower and crude oil prices. This cointegration relationship suggests a positive correlation between biodiesel and sunflower and crude oil prices. Results obtained from the parametric VECM provide evidence that biodiesel is the only variable that adjusts to deviations from the long-run equilibrium relationship. This finding is expected and is consistent with previous research (Balcombe and Rapsomanikis, 2008; Serra et al, 2011). Multivariate local linear regression shows that the speed of this adjustment is larger when biodiesel is cheap than when it is expensive. Variation in the local estimates of the VECM parameters is however small, which provides evidence that price adjustments are well represented by a linear model. Generalized impulse response functions suggest that shocks to both sunflower and crude oil prices cause changes in biodiesel prices in the same direction. Increases in energy prices are also found to increase sunflower oil price levels.

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Figure 1. Monthly price series.

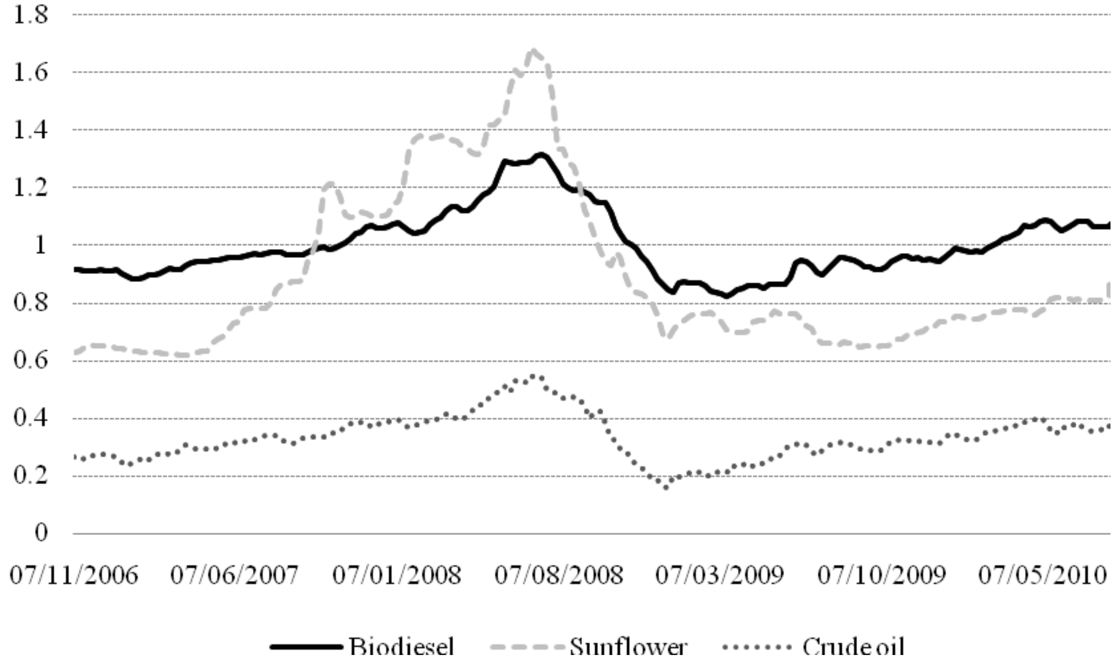


Figure 2. Distribution of localized estimates of the parameters showing adjustment to equilibrium.

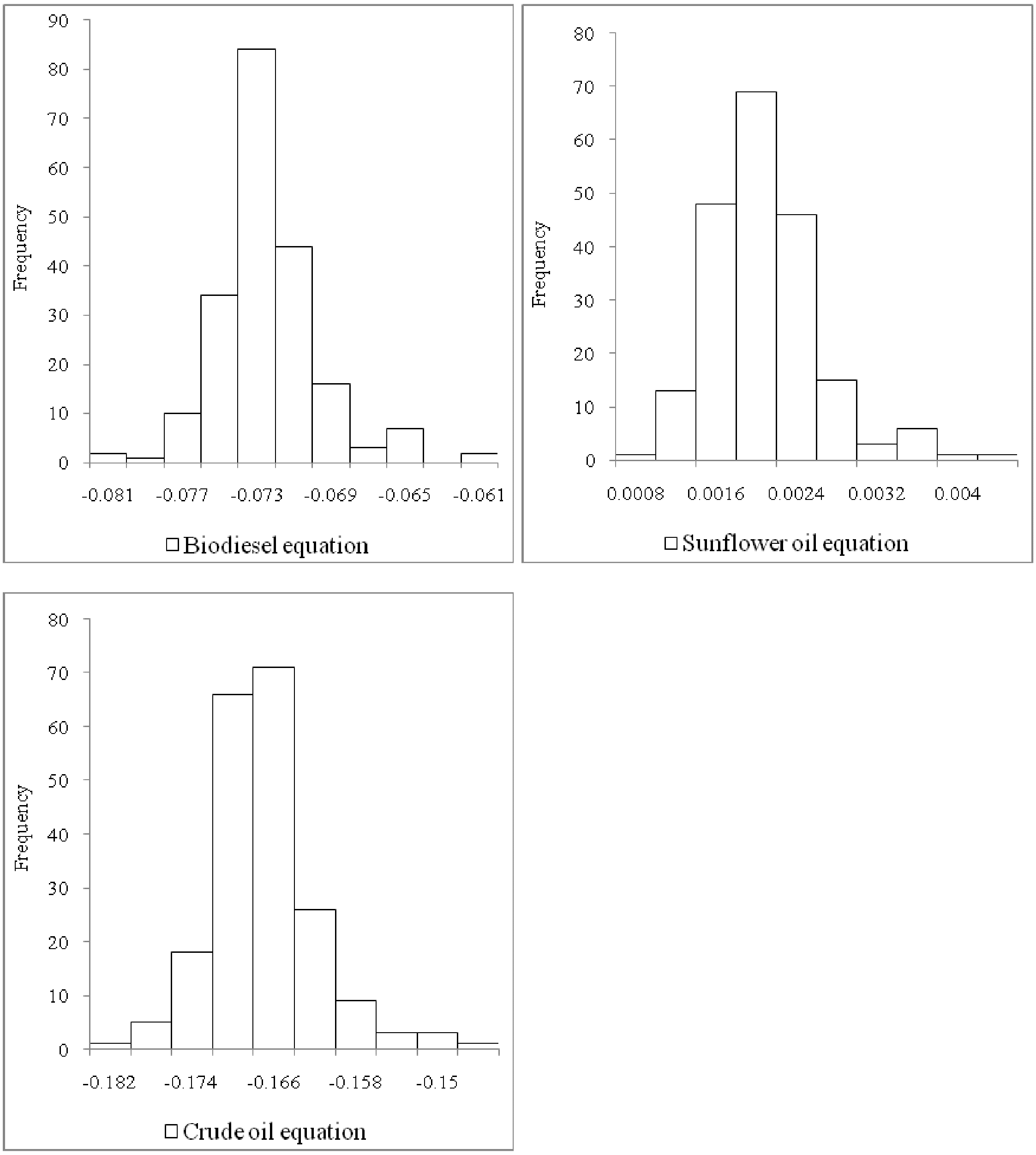
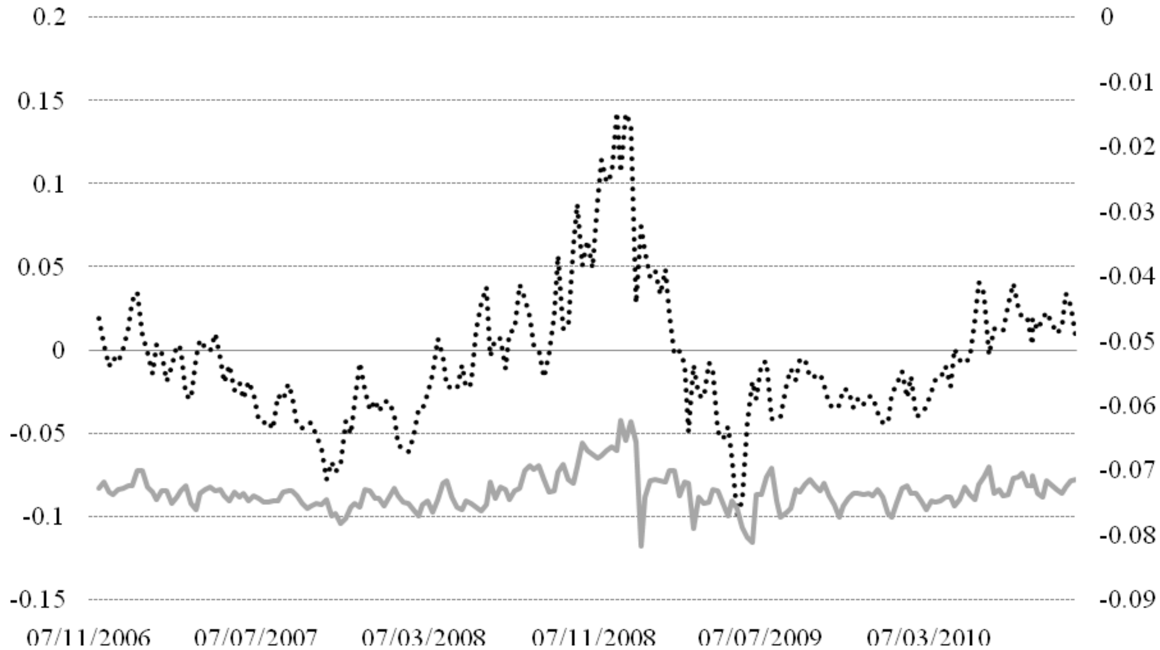


Figure 3. Evolution over time of the error correction term and the parameter showing adjustment of biodiesel price to this term.



Notes: The error correction term v_{t-1} is represented by the doted black line and is plotted on the left-hand side axis. The error correction parameter, presented by the continuous silver line, is plotted on the right-hand side axis.

Figure 4. Biodiesel response to a positive one standard deviation shock to the crude oil price.

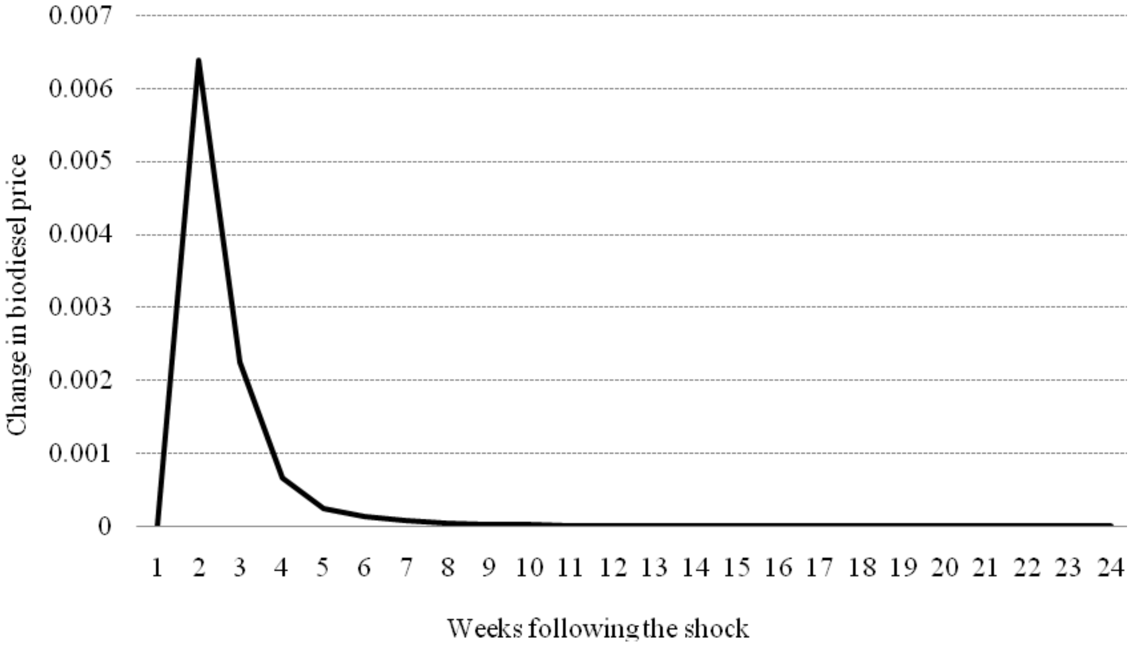


Figure 5. Biodiesel response to a positive one standard deviation shock to the sunflower oil price.

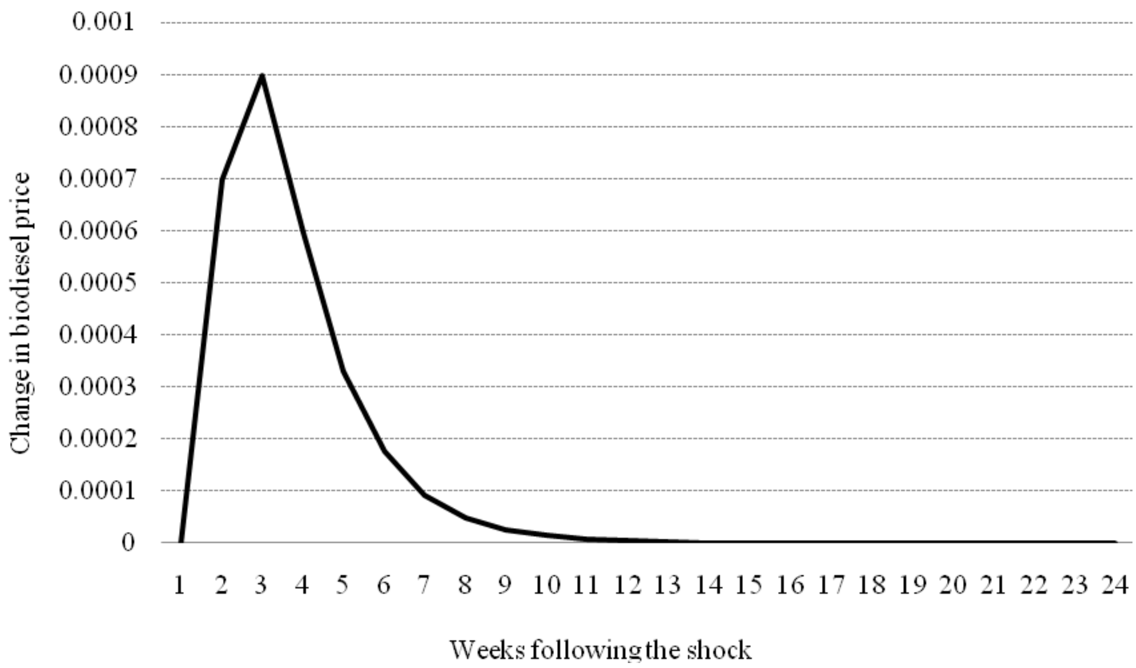


Figure 6. Sunflower oil response to a positive one standard deviation shock to the biodiesel price.

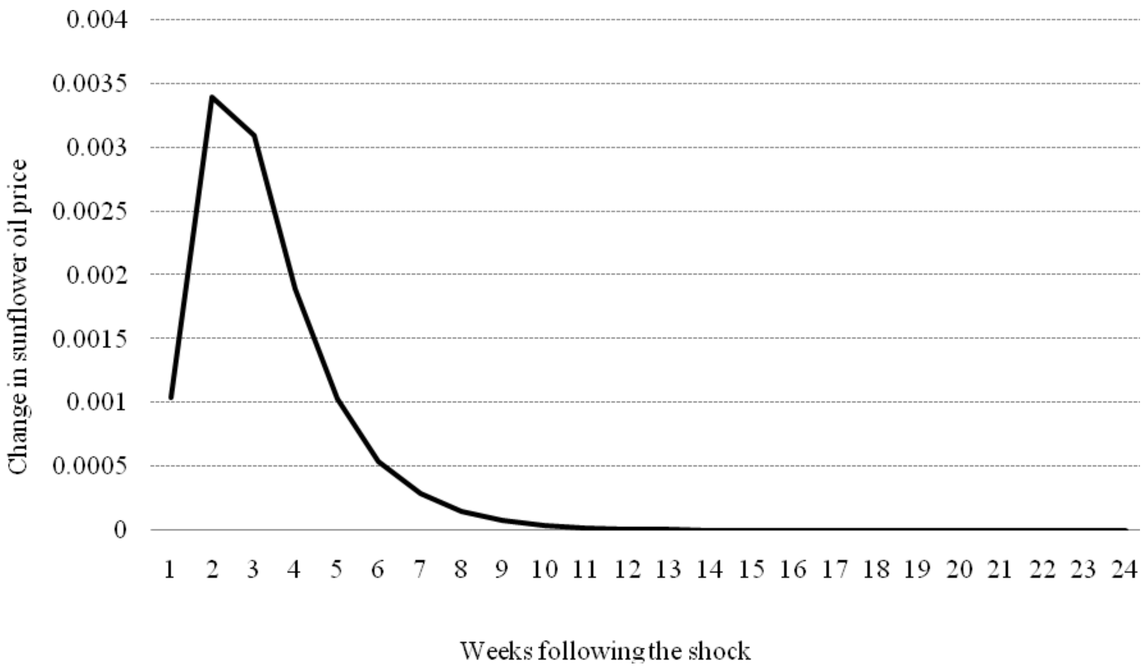


Figure 7. Sunflower oil response to a positive one standard deviation shock to the crude oil price.

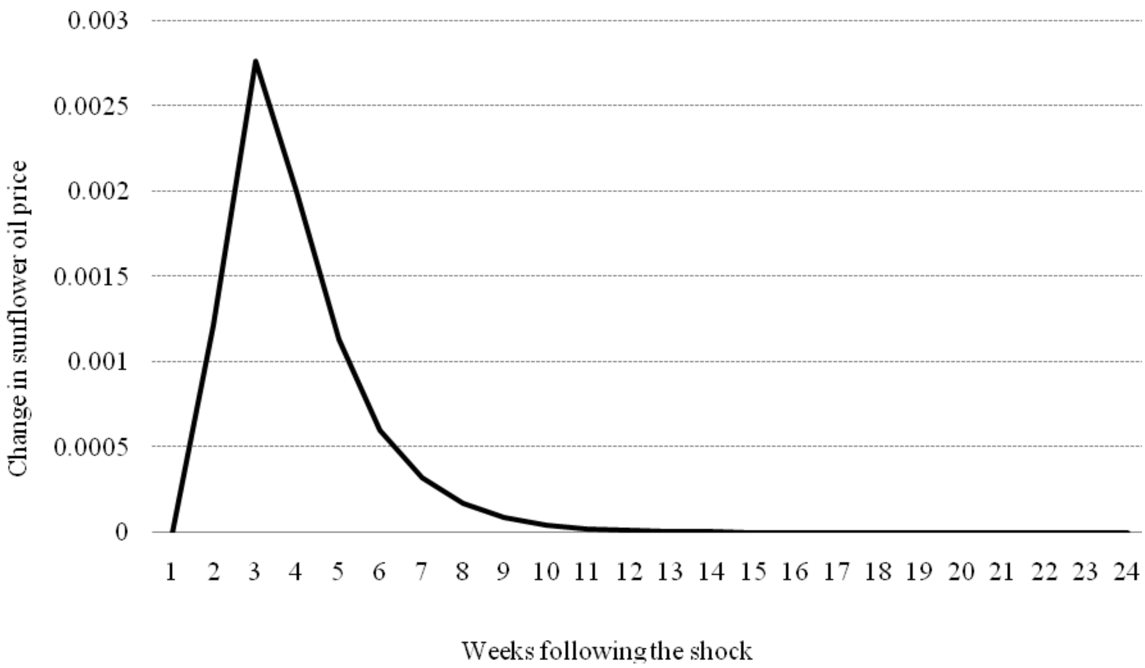


Table 1. Johansen λ_{trace} test for cointegration and cointegration relationship.

Ho	Ha	λ_{trace}	P-value
$r = 0$	$r > 0$	121.461	0.000
$r \leq 1$	$r > 1$	12.211	0.443
$r \leq 2$	$r > 2$	1.798	0.811

Cointegration relationship

(standard errors in parenthesis)

$$P_B - 0.079^{**} P_S - 0.398^{**} P_C - 0.465^{**} = Ect$$

(0.019) (0.020) (0.021)

Note: r is the cointegration rank.

** denotes statistical significance at the 5 % level.

Table 2. Estimation of the Vector Error Correction Model.

Dependent variable	Biodiesel price equation	Sunflower oil price equation	Crude oil price equation
ΔLP_{Bt-1}	0.293**(0.453)	0.335** (0.149)	-0.061(0.270)
ΔLP_{St-1}	0.000 (0.019)	0.452** (0.064)	0.171(0.115)
ΔLP_{Ct-1}	0.157**(0.014)	0.030 (0.046)	0.057(0.084)
Constant	-0.000 (0.000)	0.001 (0.002)	0.000(0.003)
v_{t-1}	-0.067**(0.017)	0.003 (0.056)	-0.167(0.102)

Notes: Numbers in parentheses are standard errors.

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

Chapter 5

Conclusions and suggestions for future works

Since the 2000s, international agricultural markets have been subject to different shocks that have caused very abrupt price changes. Market instability has been specially intense since 2006 when food prices started an increase that led to unprecedented levels between 2007 and 2008. While in the second half of 2008 prices went back to levels close to the ones before the increase, market turbulences returned on 2010 and early 2011 (FAO-OECD, 2011). According to FAO-OECD (2011) forecasts, food prices are expected to continue to be unstable in the 2010 decade. Price instability is likely to have important economic impacts both at the macro and micro-economic levels.

One of the social and political concerns that have arisen as a result of food market shocks is the extent and speed with which price shocks are passed along the marketing chain. This dissertation sheds light on this issue by focusing on food scares and the biofuels market boom. In particular, the impacts of the Bovine Spongiform Encephalopathy (BSE) crisis on price relationships among farm and retail markets for bovine in Spain have been analysed first. The effect of the Avian Influenza (AI) food scare on poultry price behavior in Egypt has been investigated in the second place. Finally, the dissertation has focused on price linkages and price transmission patterns between biodiesel, sunflower and crude oil prices in Spain.

To achieve the aforementioned objectives, the latest advances in time series econometrics have been applied. Specifically, a Regime Switching Vector Error Correction Model (RSVECM) has been estimated in order to study the Spanish beef market chain during the BSE crisis. In order to reflect consumer awareness of the BSE crisis and allow for price adjustment to depend on the magnitude of the crisis, a BSE scare index has been built on the basis of a monthly count of newspaper articles on the zoonosis published in a major Spanish newspaper and using Chern and Zuo (1997) methodology. The empirical model also utilizes two series of monthly farm-gate and retail beef prices obtained from the Spanish Ministry of Agriculture.

A Smooth Transition Vector Error Correction Model (STVECM), which is more flexible than RSVECMs, has been used to assess the second specific objective. An AI food scare index has been developed upon a monthly count of newspaper articles published in the most widely read Egyptian newspaper and used as a transition variable within the STVECM. Two series of monthly poultry prices (wholesale and consumer) have also been used to estimate the model.

Results of the food scare analyses suggest that distinct levels of the marketing chain respond differently to these crises. We find upstream prices in the marketing chain to be characterized by higher price adjustments compared to consumer prices that are sticky and slowly-responsive to food scares. The magnitude of the adjustment of upstream prices is found to depend on the magnitude of the food scare. This may increase the likelihood of upstream economic agents abandoning the sector when the magnitude of the crisis is substantially high and probably increases the need for public intervention if this is to be prevented. These results are expected and are compatible with previous research (see, for example, Livanis and Moss 2005, Lloyd et al. 2001 and Jaenicke and Reiter 2003).

Finally, error correction models estimated using multivariate local linear regression and parametric techniques are used to determine price linkages and price transmission patterns between food and energy prices in Spain. Weekly biodiesel, sunflower and crude oil prices have been used in the empirical analysis. The estimated parametric error correction model suggests that the three prices considered are linked by a long-run equilibrium relationship and that biodiesel is the only variable that adjusts to deviations from this parity. This finding is expected and is consistent with previous research (Balcombe and Rapsomanikis, 2008; Serra et al. 2011). Multivariate local linear regression shows that shocks that increase biodiesel prices will trigger faster responses than shocks that reduce biodiesel prices. Generalized impulse response functions suggest that shocks to both sunflower and crude oil prices cause changes in biodiesel prices in the same direction. An increase in energy prices results in an increase in sunflower oil price levels through short-run price dynamics.

Contributions of this dissertation to the food price transmission literature include the application of econometric techniques not yet used by previous research, the use of food scare indices in order to allow for price adjustments to depend on the magnitude of the food scare, as well as the analysis of markets not previously studied.

Considering recent forecasts pointing towards higher and more volatile commodity prices in the future (FAO-OECD, 2011) it seems that analyses that address price responses to food market shocks will continue to be relevant in future years. FAO (2008) has also emphasized that the impacts of market shocks cannot be generalized to different products and countries due to the relevance of the particularities characterizing each specific market. In this context, I would like to finish this dissertation by outlining some points and suggestions for future research. First, it would be interesting to apply the techniques proposed in this dissertation to analyze other market shocks that have not yet been studied such as the recent e-coli outbreak. This would allow assessing to what extent our results can be generalized to other settings.

Second, to deal with sudden market shocks requires not only an estimation of how price levels change, but also of how their volatility is affected (Serra, 2011). Bénabou and Gertner (1993) show that volatility in retail prices may result in an increase search costs, reducing the incentive for consumer search and increasing retailer market power. Serra (2011) also argues that volatility may trigger the exit of many small agricultural producers unable to cope with lower and much more volatile prices. Recent food price volatility has turned the political agenda onto mitigation and management of this volatility. Though the financial economics literature has paid special attention to explicit modeling of price volatility, the food economics literature has been more concerned with food price levels. The methods used in this dissertation focus on studying food price levels. Though some inferences can be drawn regarding food price volatility, this volatility is not explicitly modeled. Hence, it would be interesting to apply price volatility models to investigate the impacts of recent market shocks on food price instability. Recent FAO publications (Prakash, 2011), as well as recent EU calls (KBBE.2012.1.4-05 Volatility of agricultural commodity markets) are an indicator that research in agrofood markets in the following years will devote much attention to this issue.

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