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Vertical Price Transmission in the Italian Milk Supply Chain: Understanding the Role of Distributors, CAP Reforms, and Market (Non) Fundamentals.

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Contents

List of Figures

List of Tables

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Introduction

The analysis of marketing's margins and vertical price transmission (VPT) process along the food chain has recently attracted considerable interest among agricultural economists, mainly due to its welfare and policy implications. In recent meta-analysis research, Kouyaté and von Cramon-Taubadel (2016) spot around 500 works for which price transmission (PT) dynamics was the central topic. Being the primary mechanism by which information is conveyed, prices are the first link among market economic agents, driving both strategic and structural decisions along the supply chain (SC) (Lloyd, 2016). The last twenty years undoubtedly represented a very tumultuous run for the agrifood markets worldwide. Mergers and acquisitions (especially at the processing and retail levels), as well as vertical integration processes, have raised worries about concentration level within the agro-food industry (Sexton, 2013), with non-linear price linkages often related to market power exertion. Moreover, disentangling how recent world agricultural commodities price fluctuations have been pass-through along food chains, and in which way the CAP framed PT mechanisms in European agricultural markets, is nowadays a priority for policymakers and the academic community, allowing for a better understanding of the food sector functioning (McCorriston, 2015a). The degree and the speed of transmission reflect how much supply chain levels are integrated, providing an efficiency measure of the chain considered (Abdulai, 2002; Ben-Kaabia and Gil, 2007; von Cramon-Taubadel and Loy, 1996; Goodwin and Holt, 1999; Serra and Goodwin, 2003). The European Commission explicitly recognizes the importance of VPT analysis, stating that the "[…] adjustment of the food supply chain to price changes is an important characteristic of the functioning of markets as it reflects the nature, structure, and organization of the chain. Measuring the degree of vertical price transmission thus helps to identify potential market failures." (European Commission, 2009, p. 5). Market efficiency presupposes that in a competitive market, with perfect information, arbitrage secures that price differentials along the chain reflect the cost of marketing services, and that price shocks are instantaneously transmitted between economic agents. Nevertheless, a substantial share of works on PT concludes there is a significant gap between price dynamics and the theoretical economic background, since depending on the nature of the price shock, the mechanism of pass-through differs, leading to asymmetric responses (Ben-Kaabia and Gil, 2007; von Cramon-Taubadel, 1998; Hassouneh et al., 2012; Meyer and von Cramon-Taubadel, 2004; Serra et al., 2011; among others). Economic theory treats asymmetries as the exception, when, according to general results they are rather the rule. This generated significant interest in scholars who investigate non-linear price transmission dynamics. This thesis made use of different time series models, but all belong to what is defined as a non-structural approach. Such econometric techniques aim at identifying empirical (ir)regularities in the price links, without imposing any theoretical structure. On the one hand, critics often charge that such models fail to provide any economic insight, and are not able to distinguish price patterns under different theories (Miller and Hayenga, 2001). Accordingly, the usage of such modeling techniques have to be accompanied by a thorough description of the investigated SC, the agents participating in and the existing relationships among them (Ben-Kaabia and Gil, 2007; Schiefer and Deiters, 2015). However, because of their low-demanding volume of data, they represent an agile and cemented methodology for characterizing price behavior, shedding light on economic indicators' patterns (Serra, 2013).

My first approach to the economics and modeling of price transmission dynamics happened during my period at the Mediterranean Agronomic Institute of Zaragoza (IAMZ) as a Master student. My MSc Thesis focussed on the PT process in the Italian fluid milk supply chain using a standard linear non-structural time series model. So that, when I applied for a Ph.D. position at the University of Parma I decided to submit a proposal in which I would have centered my research on the dynamics of PT in mainly two (interrelated) markets: the dairy and the cereal sectors. Indeed, the dairy represents a fascinating sector due to its current relevance in EU policy debates. The milk market has been probably the most intervened agricultural sector under the CAP, and the final steps of a broad liberalization process began in the 2000s are becoming effective. Furthermore, dairy farmers are claiming the distribution is eroding their margins, given its high market shares regarding households' food and beverage purchases, and, hence, the exertion of market power. The interrelation with the cereal markets – being cereals and oil crops raw inputs for bovine feeding and representing more than half of the production cost for Italian producers – makes the system even more complicated and intriguing regarding price transmission dynamics. The 2007 price increase for many commodities opened up a fierce debate over the causes of such rise, raising concerns over the role played by market non-fundamentals in shaping market prices (i.e., biofuels, oil, and financialization). Therefore, I was intrigued by the complex system characterizing an agricultural commodity like fluid milk, which represents a key sector for both Italian and European agricultural economies. While trying to draw soundly-economic conclusions from my first work on PT, I found myself wandering from one study to another, gathering as much information I could, from the policy impacts to the increased volatility in the post-crisis era. So that, I decided to get more into the analysis of the sector as well as into the econometrics surrounding the PT analysis, accounting for non-linear models capturing potential asymmetries. The three chapters composing the following Ph.D. thesis deal with three macro events that have characterized the agrifood industry in the last 20 years: the concentration at the retail stage, the policy reform pushing towards a more liberalized EU agricultural market, and the role played by non-fundamentals in shaping price cycles.

The first research question I wanted to address related to the very roots of every market analysis: understanding the functioning of the SC, unveiling the nature of the existing relationships between agents, with a particular focus on the downstream levels, namely the industrial processor and the retailer. Indeed, "[…] ensuring the effectiveness and efficiency of the food supply chain is crucial to raise its competitiveness at the benefits of both consumers, with lower prices, and stakeholders of the chain, with a sustainable distribution of value added. It is also essential in order to ensure that the various actors of the economy fully benefit from agricultural policy reform towards more market orientation" (European Commission, 2009, p. 3). The dairy sector, according to Food Drink Europe, features 14% of the annual turnover of all food industries in the EU, being one of the most relevant agri-food sectors in economic terms. Italy is the seventh milk producer providing 7% of the total European cow milk output (data from Eurostat, 2017), and the dairy accounts for 12% of the total Italian agrifood turnover (Gonano and Mambriani, 2014). These numbers give a clear idea to the reader about the socioeconomic importance of milk production for both Italian and European scenarios. Furthermore, the fluid milk sector has been under the spotlight for the last ten years, with farmers' unions leading numerous public demonstrations claiming an unfair value distribution along the chain and more often to be producing at a loss. The lion share of fluid milk purchases in Europe is channeled through the so-called modern distributors (MD), and consumer food prices not always respond to input price spikes, drawing different "[…] food inflationary experiences" (McCorriston, 2015b, p. 21). This woke-up certain curiosity towards retail prices' behavior, and to what extent raw prices shocks were transmitted to final consumer prices depending on the structure of the investigated SC, with special attention paid to the level of concentration (Lloyd et al., 2015; Swinnen and Vandeplas, 2015). Indeed, there is increasing concern in the EU about the "[…] functioning of the EU food supply chain and the distribution of value-added between primary producers, food processors, wholesalers and retailers", spotting structural weaknesses in the EU agrifood SCs, such as a fragmented agricultural sector plagued by unequal bargaining powers, which contributes to a potential slow and asymmetric PT mechanism between agents (McCorriston, 2015b; McCorriston and Von Cramon-Taubadel, 2016). However, when compared to the number of research focussing on PT in world agricultural markets, the share of those trying to disentangle the nature of price dynamics in the very end of the SC is rather small (McCorriston, 2015b). The limited access to scanner price data - which may uncover retail price dynamics and building up a more general theory about price behavior – hampers the growth of retail-based PT analysis. Relying upon such price data would shed light on new market phenomena such as vertical coordination and contracts, and product differentiation. Indeed, very little has been done regarding the latter, with reported generalistic retail prices hiding part of the story, as differenced products may entail different PT mechanism (McCorriston, 2015a). Using a unique dataset of retail-scanner prices for the conventional fluid milk and its organic counterpart, I described two different mechanisms of transmission according to specific features of the SC, and, accordingly, different retailing strategies.

The second research question I discussed relates to the impact of the CAP on the Italian milk price dynamics. European CAP heavily protected the dairy sector for a very long time, via a large number of measures that have been dismantled since the early reform of the 2000s, boosting farmers strikes and protests all around the EU. The recent reforms of the Common Agricultural Policy (CAP), pushing towards the liberalization of the European agrifood market, has enhanced their disappointment. Starting in 1992 with the McSharry Reform, the CAP witnessed several interventions aiming at a dramatical reduction of the EU's protectionism and support, and causing agricultural prices experiencing much larger variations than in the past (Bonnet et al., 2015). Since the very beginning of the CAP system, milk has been provided with substantial protectionist measures (i.e., guaranteed prices, export subsidies, quotas, among others). However, is during the 80s when international but also within-EU partners started pushing for a more liberalized agricultural market that reduces distortions on the world market. That is when the milk quota system was put in place. The 90s witnessed a reduction in price supports, as well as export subsidies and tariffs, introducing payments for land and animals, according to the Uruguay Round Agreement on Agriculture under the GATT(WTO) (Swinnen et al., 2014). The 2000s probably represent the most fundamental period for the EU-dairy sector, following the so-called Luxembourg (or Fischler) Reform of 2003. The latter comprises the abandon of the quota system by March 2015, with an increase in quota thresholds in 2006, 2007, and 2008 (the "soft-landing" measure), together with a decrease in intervention price for butter and skimmed milk powder during 2004-2007 period and the elimination of export subsidies (Assefa et al., 2016; Kloosterboer, 2016). Relying upon more general price series, provided by public databases, I investigated whether policy reforms affected the PT process between Italian milk processors and retailers, accounting for structural breaks in the cointegrating relationship. Results indicate that the set of reforms impacted the nature of Italian PT process eliminating asymmetries, despite the increased volatility hampered the speed of adjustment of the market to the equilibrium.

The third research question deals with the impact market fundamentals and exogenous factors, i.e., energy prices and financial-related variables, may have on the VPT dynamics between Italian corn and dairy-cow feed prices. The agricultural price boom of 2007-2008 has been defined as the largest since the crisis of 1973-1975. The latter was characterized for being "[…] single-factorial and temporary […] food crisis with […] limited geographical scope” (Garrido et al., 2016, p. 1), whereas the 2007 crisis features rather opposite characters, that is global in its scope, multi-factorial, and sustained in time (Garrido et al., 2016). Since the sprout of the crisis, scholars tried to unveil the causes of such price rise, but a consensus has not been reached. Exogenous factors - especially energy prices and financial activity, which constitute the ‘Masters Hypothesis’ (Irwin and Sanders, 2012) - have been blamed as the crisis’ triggers by some, but discarded by others[1]. As indicated in Brümmer, Korn, Schlüßler, Jaghdani, et al. (2016), the literature seems to point at market-fundamentals as (still) the primary driver of price increases. However, as they posit, agricultural markets are, nowadays more than ever, interrelated with non-agricultural sectors. Indeed, cereals and oil crops represent feed for dairy cows (especially maize and soybeans), and more than half of operational cost for the average Italian milk producer is constituted by compound feed purchases. We then propose a new methodological approach, which allows for the non-continuous cointegrating relationship between two price series. Considering the cemented assumption of the far widely used Error Correction Model (ECM) technique stating the co-movement is continuous, we consider our approach innovative, both for its econometric and economic insights. Considering energy-related price series for both crude oil and biofuels, as well as financial-derivative variables, we conclude that non-fundamentals have negligible effects on the VPT mechanism in the Italian dairy feed SC, while market fundamentals (i.e., supply and demand) still play a relevant role in shaping price cycles.

As McCorriston (2015b) posits, price transmission matters because of three main reasons: first, it deepens the understanding of how the food inflation process is passing-through, from upstream chains down to the final consumer stage (i.e., retailers); second, it helps out in describing how a given SC works, if there exist some inefficiencies, and to formulate hipotheisis over the source of such inefficient behaviours; last, but not least, it facilitates the understanding on which agents of the SC the burden of price changes is taking place, providing a basis for policy assessment and contributing to the debate of the distributional effects in the food system. Nevertheless, according to the tremendous strain of works in literature it is clear that results from the literature suggest that are market and country-specific (García-Germán et al., 2016; Lloyd et al., 2015; McCorriston, 2015b), so that generalization of results can be misleading since each sector for each geographical area features a diverse set of economic, organizational, and societal elements shaping the relationship along the chain.

[1] See Chapter 3 of the present thesis for a more comprehensive literature review.

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1. Price Transmission Dynamics for Quality-Certified Food Products. A Comparison between Conventional and Organic Fluid Milk in Italy

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Abstract

Despite the vast amount of works investigating price transmission process in diverse agri-food markets, very little has been said about quality-differentiated products. In this paper, we propose a comparison between the conventional and the organic fluid milk sectors in Italy aiming at deepening the understanding of the economic organization and functioning of one of the essential agri-foods in Italy. Using a unique dataset featuring processor and retail (scanner) prices for the two types of milk we estimate (M)TAR models to account for asymmetric price movements in both sectors, but price transmission results symmetric eventually. The VECM estimations and IRFs analysis provide significant insights on differences between the two markets. [EconLit citations: Q130, Q110, C590]

1.1. Introduction to the Study

The analysis of price transmission (PT) mechanism along food chains has attracted considerable interest among agricultural economists, mainly due to its welfare and policy implications. Price is the first link between economic agents, driving their strategic and structural decisions and giving a good measure of integration and, thus, chain dynamics (Abdulai, 2002; Goodwin and Holt, 1999; Lloyd, 2016; Serra and Goodwin, 2003).

Despite the increasing literature on price transmission, there is a lack of studies investigating PT for quality-differentiated agricultural markets, and, therefore, a comparison with their conventional counterparts cannot be made. Consumer demand for healthier, safer, environmental-friendly and premium quality foodstuff is increasing consistently, shaping the food industry offer, with more quality-differentiated products bearing specific claims becoming available (Unnevehr et al., 2010). Among those, we focussed on the organic-labeled fluid milk. Indeed, the demand for organic food products is on the rise in Europe, with a market value estimated at 20.8 billion euros in 2012. The EU policymakers recognized the importance of this emerging market through supporting organic farming, in light of its contribution in producing public goods (Meredith et al., 2014). Furthermore, Italy represents the fourth organic market in Europe with more than 2 billion Euros of retail sales, and the second one regarding both organic producers, with around 52 thousand operators, and organic land, with more than 1.4 million hectares in 2015 (Willer et al., 2017).

The objective of this study and its main contribution to the existing literature is the discovering of the price transmission mechanism concerning an organic food product, providing a comparison to its conventional counterpart, and deepening the understanding of the economic organization and functioning of both markets. The Italian fluid milk market has been chosen as the case study as it represents one of the most important agro-food sectors in Italy. The dairy industry is a major asset which is worth 15 billion euros, 11.5% of the food industry turnover, and with an employed workforce of about 30,000 people (Gonano and Mambriani, 2014). Moreover, Italy is the 7th EU-28 milk producer, providing 7% of total EU (cow) milk (EUROSTAT, 2017). Hence, the conventional fluid milk market represents the benchmark for comparison with organic price transmission dynamics.

A unique dataset has been used for the analysis, featuring processor and retail (scanner) prices for the two types of milk, provided by one of the most prominent Italian food distributors. This gives insights into retailer marketing strategies for conventional and differentiated (i.e., organic) products. Indeed, price transmission studies largely use data for aggregate categories (e.g., dairy, meat, apples), an approach which prevents generating robust results when the research objective concerns quality-differentiated products. Time series econometrics has been applied for investigating price transmission dynamics. The potential role of asymmetries is then explored through nonlinear univariate models. When linearity tests accept the null of linearity, a Vector Error Correction Model (VECM) is used to interpret long-run relationships, and Impulse Response Functions (IRFs) are used to understand short-run dynamics.

The work is organized as follows. Section 2 provides a brief analysis of the organic and the conventional fluid milk sectors, while Section 3 illustrates the theoretical framework. Section 1.4 describes the type of data used and the applied methodological approach. Section 1.5 shows the empirical results for each system, whereas Section 1.6 provides a discussion of the results and Section 1.7 conclude.

1.2. The Organic and Conventional Fluid Milk Sectors in Italy

The Italian dairy sector is undergoing a profound crisis, and organic production has grown tremendously as an instrument of differentiation and, thus, margin stabilization, representing a sort of safety-net for troubled milk producers. Private dairy companies transform the raw milk delivered by farmers, producing 92% of the total fluid milk (ISTAT, 2016), which reaches consumers through retailers. In fact, modern distribution (MD) is currently the primary retail format in Italy, consisting of all those points of sale having large surfaces and self-service (AGCM, 2013) (see also Sckokai (2014) for a detailed classification of MD channels). According to the most recent data available, the CR(5) (Concentration ratio of the first five companies) of the MD reaches almost 50%, with nearly 90% of household spending on fluid milk concentrated on this channel, justifying the concerns expressed by society and institutions over its impact on chain performance (Federdistribuzione, 2013). On the supply side, raw conventional milk production relies upon 157 thousand dairy farms (with at least one bovine head), accruing for more than 4.6 million of heads, and producing more than 11 million of tonnes of raw milk. Eventually, the total amount of fluid milk produced reaches 2.3 million of tonnes in 2015 (ISTAT), with a sales value around 2.5 billion euros (Gonano and Mambriani, 2014). When compared to the numbers of the organic sector, the latter still represents a niche market within the Italian milk industry. Relying on the statistics provided by the Official Italian National Accreditation Body (ACCREDIA), there are 255 organic dairy farms currently active on the Italian territory, and more than 57 thousand dairy cows. The total organic raw milk production achieved in 2015 was 215 thousand tonnes, of which 23.5 thousand tonnes destined to fluid milk (statistics from SINAB - National Information System on Organic Farming). According to the Organic Data Network provided by Willer (2015), dairy products accrue for 18% of total organic sales, around 77 million euros, after fruits and vegetables (27%). Within the category of dairies, fluid milk represents the second most-sold product after yogurts, with 34% of overall sales, around 27 million euros. Organic milk represents almost 2% of total milk sales and 3.5% of total milk production in Italy. Organic supply chains are usually regarded as short supply chains, that is, the number of actors involved, from production to retailing, is lower than in conventional agrifood chains. Upstream agents seem to be more organized, since the constitution of producers' organizations (POs) has been widely promoted and boosted by the European Commission and, consequently, by the national government. Moreover, the distribution sector is more competitive, since the MD represents just 27.4% of total organic sales, with specialized organic shops accounting for the lion's share (45%) (Romeo and Bteich, 2014). Accordingly, one might expect in the organic market that retailers have a lower market power, if any, as the supply is also quite concentrated. As a consequence, we would expect symmetric responses from processors and retailers after unanticipated supply and demand shocks. In relation to the magnitude and the length of responses, one should expect longer responses as the marketing channels are less concentrated than in the conventional fluid milk supply chain

1.3. Theoretical Framework

In this section, we provide a simple theoretical framework to explain which type of relations exist between the industrial processors of milk and the food distributors in terms of prices. Within each of the two chains, the product is considered homogeneous, produced on the Italian territory only – since it is a high-perishable product. Thus, we assumed constant returns to scale[2], and fixed-proportion technology for both agents concerned by the study

[2] See the seminal paper of McCorriston et al. (2001) for the interaction between returns to scale and market power.

(Sexton and Zhang, 2001). Given a general (inverse) consumer demand function for the fluid milk at the retail stage as

$$P_r = D(Q_r, X_r) \tag{1}$$

where Q_r is the quantity of product at the retail level, X_r is an exogenous demand shifter, and P_r is the price faced by consumer on the final market. Likewise, the (inverse) supply function of the industrial processor is defined as

$$P_p = S(Q_p, X_p) \tag{2}$$

where Q_p is the quantity of milk delivered by farmers and X_p is an exogenous supply shifter. We assumed both milk processors and retailers use a constant-return technology and that no substitution is permitted between the raw milk delivered to processors and the other input used in producing the final products. Following Verreth et al. (2015), we assumed that industrial processors act as price leader, hence retailers take the industrial processor price as given. Consequently, industrial processors determines the quantity of milk to be processed, relying on retailer's behaviour. Therefore, the retailer's profit is maximized as

$$\max_{q_i^r} \pi_i^r = D(Q_r)q_i^r - S(Q_p)q_i^p - c_i^r q_i^r, \tag{3}$$

where $Q_r = \sum_{i=1}^{n} q_i^r$, and c_i^r represents the costs sustained by the retailer in handling one unit of product. The first order condition of (3) is then given by[3]

$$\frac{\delta \pi_i^r}{\delta q_i^r} = P_r + q_i^r P_r'(Q_r)\frac{\delta Q_r}{\delta q_i^r} - P_p - c_i^r = 0, \tag{4}$$

from which we can derive the elasticity notation as

$$P_r\left(1 - \frac{\theta_r}{\eta_r}\right) = P_p + c_r \Rightarrow P_p(Q_r | \eta_r, \theta_r, c_r), \tag{5}$$

where $\theta_r = \frac{\delta Q_t^r}{\delta q_{i,t}^r}\frac{q_{i,t}^r}{Q_t^r}$ defines the conjectural elasticity of retailers, i.e., the degree of oligopoly power exerted at this level of the chain, while $\eta_r = \frac{\delta Q_r}{\delta P_r}\frac{P_r}{Q}$ represents the absolute value of the demand's price elasticity. The conjectural variation elasticity describes the exertion of market power by retailers on consumers, taking a value of 1 when the distributor acts as a monopolist, and a value of 0 when the market is perfectly competitive (Achayra et al., 2011; Huang and Sexton, 1996; Sexton et al., 2007). Likewise, the profit function for industrial processor is defined as

$$\max_{q_i^p} \pi_i^p = P_p(Q_r | \eta_r, \theta_r, c_r)q_i^p - c_i^p q_i^p - P_f q_i^f - C_p, \tag{6}$$

[3] See Lloyd et al. (2006) and Sexton et al. (2007) for a detailed derivation process.

where c_i^p defines the handling cost of the single processor for handling a unit of product sold, C_p defines the cost of industrial processing of the raw material q_i^f, assumed to be fixed, and the suffix f refers to the farm-gate level. According to Huang and Sexton (1996), as quoted in Verreth et al. (2015), the first order condition of (6) can be expressed as

$$P_p\left(1-\frac{\theta_p}{\eta_p}\right)=P_f+c_i^p+C_p, \qquad (7)$$

where, as in (5), $\theta_p=\frac{\delta Q_t^p}{\delta q_{i,t}^p}\frac{q_{i,t}^p}{Q_t^p}$ is the wholesaler's conjectural elasticity and $\eta_r=\frac{\delta Q_p}{\delta P_p}\frac{P_p}{Q_p}$ determines the absolute price elasticity of industrial processor demand. We then assume constant demand elasticities as well as equal conjectural variation elasticities, and the market power terms $\frac{\theta_r}{\eta_r}$ and $\frac{\theta_p}{\eta_p}$ are treated as constants (i.e., unknown parameters) and bounded $0<\frac{\theta_i}{\eta_i}<1$ where $i=r,p$. In a competitive market, $\theta_i=0$, hence the $P_r=P_p+c_r$ in (5), and likewise $P_p=P_f+c_i^p$ in (7). In sum, the retail price equal the processor price plus a constant cost, whereas the industrial processor price equals the farm-gate price plus both the handling and transformation costs. Concerning the price transmission elasticities, whenever one assumes fixed-proportion technology, the former equals the processor level's price elasticity of demand divided by the retail level's price elasticity of demand (Verreth et al., 2015):

$$\varepsilon_r=\frac{\delta P_r}{\delta P_p}\frac{P_p}{P_r}=\frac{P_r-\left(c_i^r+C_p/\left(1-\frac{\theta_r}{\eta_r}\right)\right)}{P_r}, \qquad (8)$$

where ε_r is the price transmission elasticity of retailers. A similar equation can be obtained concerning the industrial processor, but taking into account the farm-gate price of raw milk that we would not do it in the present study. In (9), whenever $\frac{\theta_r}{\eta_r}$ approaches zero, the $\varepsilon_r=\frac{P_p}{P_r}=\frac{(P_r-C_p-c_i^r)}{P_r}$, that is the share of industrial producer price in the retail price. On the contrary, if market power term approaches the unity and ε_r approaches zero, the retailer absorbs (almost) completely changes occurring at the processing level, exerting market power. In such a dramatic case, the retail price would not respond to changes at the industrial level. If one tries to understand the impact of both demand and supply shocks on both retail and processor prices, it can be specified as

$$\frac{dQ_r}{dX_r}=\frac{\partial Q_r}{\partial X_r};\ \frac{dQ_p}{dX_p}=\frac{\partial Q_p}{\partial X_p}\ . \qquad (9)$$

Of course, as Lloyd et al. (2006) pointed out, the signs of the only element on the right-hand side will depend on the shock. A decrease (increase) in fluid milk demand would entail a decrease (increase) in the price of the product, and the term will have a negative (positive) sign. Likewise, a positive (negative) shock on the supply side would mean a lower (higher) price on the processor side. The previous framework, however, assumes price transmission symmetry, as most of the theoretical literature. However, most of the recent empirical work has shown that asymmetries in price transmission (APT) are the general rule and not the exception. Peltzman (2000) was the first generalizing this idea, investigating a large number of markets and finding asymmetric behaviors in more than two of every three analyzed sectors. Indeed, a significant strain of literature found APT, and several hypotheses have been explored by academics to explain such dynamics[4]. Retailers' market power is the most frequently identified cause (see among others Bailey and Brorsen, 1989; Verreth et al., 2015; Borenstein et al., 1997; Lloyd et al., 2006; McCorriston et al., 2001; Sckokai et al., 2013; Sexton, 2013; Shrinivas and Gómez, 2016; Simioni et al., 2013), although further studies (Acosta and Valdés, 2014; Bettendorf and Verboven, 2000; Peltzman, 2000; Serra and Goodwin, 2003) conclude the exertion of market power and high concentration ratios do not thoroughly match with asymmetries. Additional causes of APT have been debated, including substitutability between agricultural and other marketing inputs (Bettendorf and Verboven, 2000; McCorriston et al., 1998), adjustment costs (Azzam, 1999; Chavas and Mehta, 2004), product perishability (Santeramo, 2015; Santeramo and von Cramon-Taubadel, 2016; Ward, 1982), policy intervention (Brümmer et al., 2009; Cacchiarelli et al., 2016; Esposti and Listorti, 2013a; Ihle et al., 2012; Kinnucan and Forker, 1987; Lee and Gómez, 2013; Santeramo and Cioffi, 2012), asymmetric information (Bailey and Brorsen, 1989) and inventory costs (Reagan and Weitzman, 1982). Tifaoui and von Cramon-Taubadel (2016) recently investigated the impact of the temporary sale price (TSP) on price transmission for butter in Germany, arguing the "valleys" generated from TSP increase the speed and asymmetry of vertical price transmission. Recently, McLaren (2015) developed a theoretical framework to explain the existence of asymmetric price dynamics if there exist sufficiently convex marginal cost curves for market intermediaries, consistent with the monopsony power exerted by them. Investigating the relation between export intermediaries and farmers for a broad range of agricultural products and countries, he found that decreases in farm price are far more completely passed to FOB prices than increases. Therefore, intermediaries use their market power to benefit from stretching margins, since they re-sell the agricultural product on the international market. Certainly, this logic can be applied to other types of market linkages between different agents: in our specific case, between food processors and retailers. One may expect the latter exerting market power, and the so-called "rocket and feather" dynamics to take place (i.e., food-processor price increases are more fully and quickly transmitted to retail prices). According

[4] Meyer and von Cramon-Taubadel (2004) provide a thorough summary of the leading causes of asymmetry.

to McLaren (2015), the relation between retailer and industrial processor at the optimum can be specified as

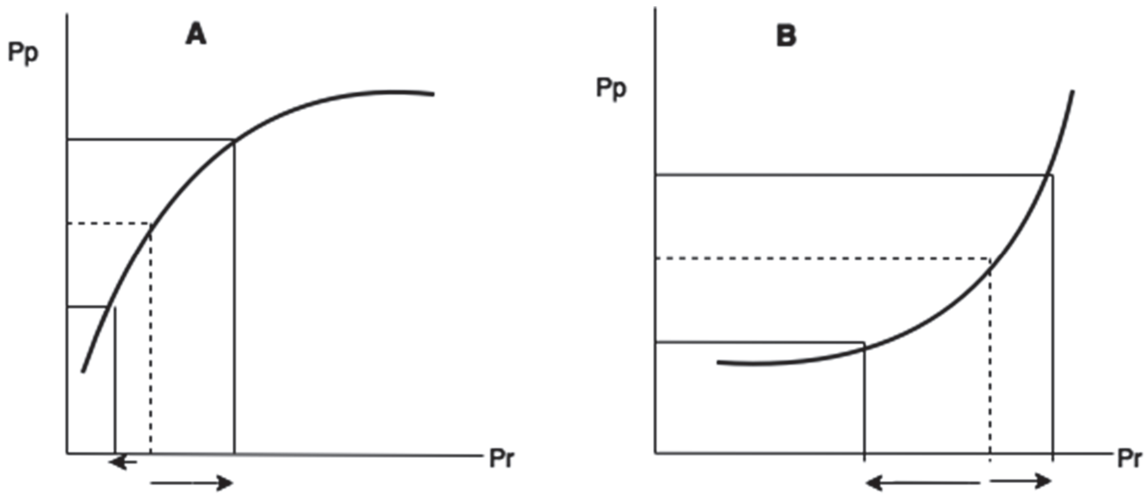
$$\overline{P_r} = w(\bar{Q}) = w\big[\bar{Q}(P_p)\big], \tag{10}$$

where Q is the aggregate supply of processors to retailers, being $w(\bar{Q})$ the inverse supply function of retailers and $\overline{P_r}, \bar{Q}$ the retail price and quantity maximizing the profit, respectively (for the detailed methodology and derivation see McLaren (2015)). Though, one can evaluate the presence of asymmetry taking the second derivative with respect to P_p, such as

$$\frac{d^2\overline{P_r}}{\left(dP_p\right)^2}\,. \tag{11}$$

If the latter is equal to zero, the transmission is symmetric. Restrictions were put on both supply functions' shapes, resulting in (10) to have a negative sign and, hence, producing a concave relationship between the two levels of the chain. Indeed, given the high-perishable nature of fluid milk, one should expect negative asymmetries[5] to take place, with retailer fearing spoilage (i.e., a 1% decrease in processor price, more fully passed to the retail price, Ps, than an increase) (see diagram A in Figure 1). However, a different assumption on the marginal cost curve may be set, since one rather expects positive instead of negative asymmetries. The marginal cost curve would become steeper, losing convexity and modifying the relationship between the two prices, which turns convex (see diagram B in Figure 1).

Figure 1 – The Relationship between Processor and Retailer Prices



Source: Authors' elaboration based on McLaren (2015)

[5] According to Meyer and von Cramon-Taubadel (2004), positive asymmetries are defined whenever the system responds faster to squeezing rather than stretching-margin situation. Negative asymmetries occur when the opposite situation takes place.

Most recent studies on dairy and fluid milk markets have given controversial results. In analyzing dairy products in Austria, Amador et al. (2010), demonstrate the existence of positive asymmetries between producer and retail levels, characterized by a higher degree of concentration. The same results were obtained in Capps and Sherwell (2007) and Zeng and Gould (2016), in their studies on fluid milk market in the US as well as in Bakucs et al. (2012) for what concerns Poland. On the other hand, Acosta and Valdés (2014) found that negative asymmetries exist in the Panama dairy sector, which experienced an increased concentration level in recent years. Likewise, Awokuse and Wang (2009) results of milk price transmission in the U.S. led to negative asymmetries. Serra and Goodwin (2003) found symmetric adjustment in high-perishable dairy products in Spain, despite a highly concentrated retail level. Concerning the Italian fluid milk sector, Cavicchioli (2013) and later Madau et al. (2016) detected the exertion of market power from retailing toward farmers along the fluid milk supply chain. The results obtained by Sckokai et al. (2013) in their study on Parmigiano Reggiano and Grana Padano cheeses in Italy (the two major products of the quality-certified dairy industry) find no evidence of market power from retailer toward processors, while there are signs of market power toward consumers. Table 1 resumes the results obtained from the most recent works on price transmission concerning the fluid milk market.

Table 1- Summary of Most Recent Studies on Price Transmission Mechanism in Fluid Milk Markets

Author(s)	Country	Relationship	Data Frequency	Causality	Main Results	Type of Relationship	α estimates[(a)]
(Amador et al., 2010)	Austria	Farmer-Retailer	Monthly	Farmer->Retailer	Pass-through rate more than proportional, no long-run price homogeneity. Shocks in farmer price entail persistent effects on the consumer price	Positive Asymmetries	0.02
(Capps and Sherwell, 2007)	U.S. (7 Cities)	Farmer-Retailer	Monthly	Farmer->Retailer	Decreasing farm prices: from 1 to 6 months for a full shock transmission; increasing farm prices: from 1 to 3 months	Positive Asymmetries	0.27 (+); 0.39(-)
(Zeng and Gould, 2016)	U.S. (11 Cities)	Farmer-Retailer	Monthly	Farmer->Retailer (in 6 out of 11 cities)	In the short-run, retail price responds differently depending on the sign of the change. In the long-run, it takes several months for the retail price to reflects the farmer price change	Positive Asymmetries	0.14
(Bakucs et al., 2012)	Poland and Hungary	Farmer-Retailer	Monthly	Farmer->Retailer (Hungary); Retailer->Farmer (Poland);	Different market structures entail different PT dynamics	Positive Asymmetries (Poland) Symmetric (Hungary)	0.22 0.20
(Awokuse and Wang, 2009)	U.S.	Farmer-Retailer	Monthly	Farmer <-> Retailer	The presence of asymmetric dynamics should be taken into serious consideration when designing new policy reforms	Negative Asymmetries	0.31(-); 0.04(+)
(Acosta and Valdés, 2014)	Panama	Farmer-Wholesaler	Monthly	Farmer->Wholesaler	The speed of convergence is "moderately slow" (i.e., 9% of the shock is transmitted for each period	Negative Asymmetries	0.09; 0.32(-)
(Serra and Goodwin, 2003)	Spain	Farmer-Retailer	Weekly and Monthly	Farmer->Retailer (weekly freq.); farmer prices more elastic retail prices changes (monthly freq.)	Symmetric PT in highly perishable dairy product prices. Slow adjustment of retail prices to farm price shocks	Symmetric (for high perishable dairies); Asymmetric (for long shelf-life dairies)	-

(a) Absolute values. When displayed, '+' and '-' signs refer to increasing and decreasing margins, respectively. Otherwise, the coefficient refers to linear VECM estimation. We reported only statistically significant coefficients

Source: Authors' personal elaboration

1.4. Modeling Asymmetric Price Transmission

The presence of cost frictions may entail thresholds in the price transmission mechanism (Ben-Kaabia and Gil, 2008; Lee and Gómez, 2013; Meyer, 2004), and only when deviations exceed the threshold(s) the adjustment is triggered (Abdulai, 2002; S. N. Balke and Fomby, 1997; Goodwin and Holt, 1999; Meyer and von Cramon-Taubadel, 2004). Different market structures and products' characteristics involve different search and information costs for consumers. For these reasons, threshold models may provide a more suitable modeling technique to account for these types of transaction costs. The Threshold Autoregressive (TAR) model was first discussed by Tong (1983) and later revisited by Enders and Granger (1998), who introduced a model variation, the Momentum-TAR (M-TAR), increasing their popularity since then (for recent research in agricultural economics applying a TAR framework see Abdulai (2002), Lee and Gómez (2013), Simioni et al. (2013), Goychuk and Meyers (2014), Tekgüç (2013), Awokuse and Wang (2009), Surathkal et al. (2014) and Han et al. (2016) among others). Engle and Granger (1987) showed that when two variables are co-integrated, an Error Correction Model (ECM) can be specified as

$$\Delta p_{1t} = \Phi_j(p_{2,t\text{-}1}\text{-}\beta_0\text{-}\beta_1 p_{2,t\text{-}1}) + \sum_{i=1}^{k\text{-}1} \Gamma_i \, \Delta p_{1,t\text{-}i} + \sum_{i=1}^{k\text{-}1} \delta_i \, \Delta p_{2,t\text{-}i} + \varepsilon_t, \tag{12}$$

where $p_{i,t}, i = 1,2$ are prices at two different levels of the supply chain, β_0 is a constant term, the term inside the brackets specifies the error correction mechanism (i.e., ECT), Γ_i and δ_i are matrices of short-run parameters estimating the effect of shocks on $\Delta p_{i,t}$, and ε_t is a disturbance term $\text{i.i.d.} \sim WN(0, \sigma^2)$. Concerning the TAR approach, Enders and Granger (1998) specified an alternative error correction specification, arguing that in the presence of asymmetries, the two-steps Engle and Granger approach was misspecified. Therefore, they specified the error correction term as

$$\Delta\widehat{\mu}_t = \begin{cases} \rho_1 \hat{\mu}_{t\text{-}1} + \varepsilon_t \text{ if } \hat{\mu}_{t\text{-}1} \geq \tau \\ \rho_2 \hat{\mu}_{t\text{-}1} + \varepsilon_t \text{ if } \hat{\mu}_{t\text{-}1} < \tau \end{cases}, \tag{13}$$

where $\Delta\widehat{\mu}_t$ is the differenced ECT, τ represents the threshold value (i.e., zero in this case), and where a necessary condition for $\{\hat{\mu}_t\}$ to be stationary is that $-2 < (\rho_1, \rho_2) < 0$. A formal way to quantify the adjustment process is:

$$\Delta\widehat{\mu}_t = I_t p_1 \hat{\mu}_{t\text{-}1} + (1\text{-}I_t) p_2 \hat{\mu}_{t\text{-}1} + \varepsilon_t, \tag{14}$$

where I_t is the Heaviside indicator function $I_t = \begin{cases} 1 & \text{if } \hat{\mu}_{t\text{-}1} \geq \tau \\ 0 & \text{if } \hat{\mu}_{t\text{-}1} < \tau \end{cases}$. Whenever the system is convergent $\hat{\mu}_t = 0$, whereas when $\hat{\mu}_{t-1}$ is above (below) its equilibrium value, the adjustment is $p_1\hat{\mu}_{t-1}$ ($p_2\hat{\mu}_{t-1}$). Accordingly, the error correction representation in (12) can be specified as

$$\Delta p_{1t} = I_t p_1 \hat{\mu}_{t-1} + (1 - I_t) p_2 \hat{\mu}_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta p_{1,t-i} + \sum_{i=1}^{k-1} \delta_i \, \Delta p_{2,t-i} + \varepsilon_t. \tag{15}$$

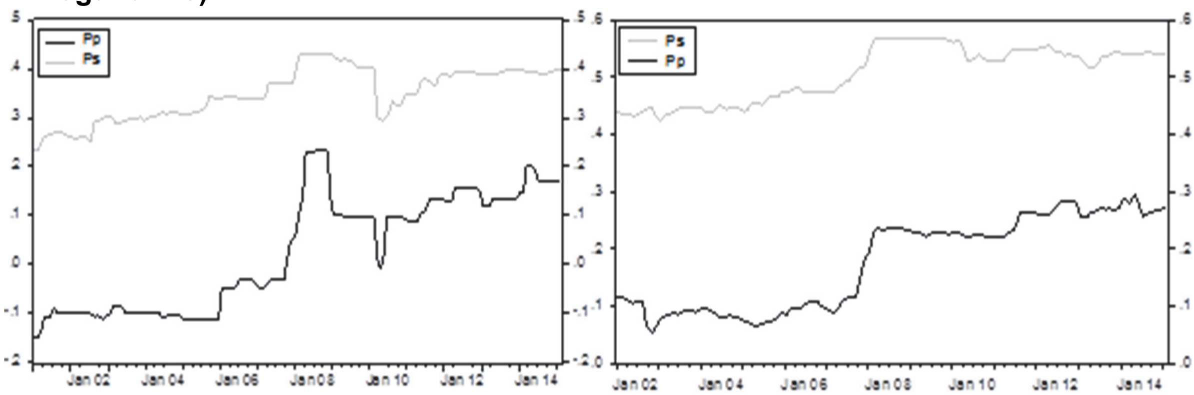
The MTAR specification described in Enders and Granger (1998), consists of adding lagged changes to the $\{\hat{\mu}_t\}$ process, such that Equation (14) can be now expressed as $\Delta\widehat{\mu_t} = I_t p_1 \hat{\mu}_{t-1} + (1 - I_t) p_2 \hat{\mu}_{t-1} + \sum_{i=1}^{k-1} \gamma_i \Delta\hat{\mu}_{t-1} + \varepsilon_t$. Accordingly, the Heaviside indicator can be now specified as $I_t = \begin{cases} 1 & if\ \Delta\hat{\mu}_{t-1} \geq 0 \\ 0 & if\ \Delta\hat{\mu}_{t-1} < 0 \end{cases}$.

The latter model is preferred when series exhibits more "momentum" in one direction. Regarding the estimation of the threshold, there is no theoretical-based argument in assuming a priori zero-threshold (Simioni et al., 2013), hence Chan's (1993) technique, which estimates a super-consistent threshold is applied. However, it is first necessary to test the existence of threshold cointegration (i.e., $H_0: \rho_1 = \rho_2 = 0$), and only when the null hypothesis is rejected in favour of threshold cointegration testing for the asymmetric adjustment (i.e., $H_0: \rho_1 = \rho_2$).

1.5. An Application to the Italian Conventional and Organic Fluid Milk System

The dataset for each conventional and organic market is composed of 170 monthly observations, covering the period 2001-2015 (see Figure 2). Prices are expressed in €/litre, defining the amount paid (per unit of product) by retailer to processor (purchasing or processor price, P_p) and the price paid by consumer on the final market (selling or consumer price, P_s). Series were transformed into their logarithmic forms, mitigating the fluctuations and increasing the likelihood of stationarity after first differencing (Hamilton, 1994). From an economic point of view, this also makes it possible to interpret results in percentage change terms (Ben-Kaabia & Gil, 2007). Unit-root tests were used to interpret the stochastic properties of the series, for identifying non-stationary series and selecting the right determinist term(s) for accurate modeling.

Figure 2 – The Organic (left) and Conventional (right) Milk Systems (prices expressed in logarithms)



Source: Authors' elaboration

After excluding the presence of a trend, a constant was included when testing for unit-root[6], since this improves the stability of results and, particularly for price transmission analysis, accounts for the current margin between prices over time. The Phillips-Perron (PP) (Phillips and Perron, 1988), the Generalised Least Square-Augmented Dickey Fuller (GLS-ADF) (Elliott et al., 1996) and the KPSS (Kwiatkowski et al., 1992) tests for unit-root were applied[7]. All the unit-root tests conducted on the two price systems led to the conclusion that the series are $I(1)$ (see Table 2). When the same battery of unit-root tests is applied to variables in first differences, they proved to be $I(0)$, confirming that the series are (stochastic) difference stationary processes.

Table 2 - Unit-Root Tests

Conventional Milk System			**Organic Milk System**		
ADF-GLS Unit Root			**ADF-GLS Unit Root**		
With Constant	(P_p)	(P_s)	With Constant	(P_p)	(P_s)
Tau	-0.194	-0.235	T-Stat	-0.0757	-0.316
P-Value	0.616	0.601	P-value	0.657	0.572
MAIC*	1	1	MAIC*	2	2
PP Test			**PP Test**		
Z_t	-1.321	-2.108	Z_t	-0.557	-1.297
P-Value	0.619	0.242	P-Value	0.876	0.630
KPSS Test			**KPSS Test**		
T-Stat	7.095	5.776	T-Stat	4.827	4.078
P-Value	<0.01	<0.01	P-value	<0.01	<0.01
Lag Truncation	1	1	Lag Truncation	2	2

*Modified Akaike Information Criterion, see Ng and Perron (2001) for more details.

Source: Authors' elaboration

1.5.1. Long-Run Relationships

Since all prices are $I(1)$, the following step consists of testing for cointegration for each pair of prices within each system. Regarding the organic market, to appropriately model price transmission, four impulse dummy variables were included, accounting for atypical spikes, taking the value of 1 for October 2007, February 2008, January 2013 and June 2014 and 0 otherwise. Likewise, due to a break in the mean of the series in the period March 2005 - May 2012, a shift dummy was included. Bearing in mind the data refers to a single private economic agent, atypical behaviors along the series could have been triggered by economic events caused by private agreements between the parties involved, preventing complete understanding of series' dynamics. The AIC (Akaike Information Criterion) suggested a

[6] The F test statistics for linear restrictions: $\beta = 0, \gamma = 0$ in the ADF regression $\Delta\Upsilon_t = \alpha + \Upsilon_{t-1} + \beta t + \sum_{j=1}^{p} \gamma_j \Delta\,\Upsilon_{t-j} + \varepsilon_t$ were 3.62 and 3.68 for conventional and organic milk, respectively, which are less than the C.V. at 5% 6.49 (Φ_3). See Dickey and Fuller (1981) for further details.

[7] The standard ADF (Dickey and Fuller, 1979, 1981) and the PP tests suffer from low power and size distortions, leading to over-reject the hypothesis of unit root (DeJong et al., 1992; Ng and Perron, 2001; Schwert, 1989). In order to overcome such limitations and provide reliable results, we use the KPSS and the GLS-ADF jointly with the PP tests.

VAR(7), and the Johansen (1988) cointegration test was applied accordingly (see Table 3), restricting both the constant and the shift dummy to the cointegration space:

$$\Delta \mathrm{P_t} = \alpha\left(\beta' \mathrm{P_{t-1}} + \mu_1' + D_{s,t}\right) + \sum_{\mathrm{i=1}}^{\mathrm{k-1}} \Gamma_{\mathrm{i}} \Delta \mathrm{P_{t-i}} + D_{i,1} + D_{i,2} + D_{i,2} + D_{i,3} + D_{i,4} + \varepsilon_{\mathrm{t}}, \qquad (16)$$

where μ_1' and $D_{s,t}$ are deterministic terms restricted to the cointegration space, namely the constant and the shift dummy respectively, the $D_{i,\mathrm{m}}, m = 1,2,3,4$ are dummy variables, as described above, α is the so-called loading matrix and β' is the long-run coefficient.

Table 3 – Organic Milk System: Johansen Co-integration test

$p-r$ [a]	r [b]	λ_{trace}	C.V.	p-Value	λ_{max}	C.V.	p-Value
2	0	22.362	20.164	0.023	14.09	15.892	0.094
1	1	8.272	9.142	0.074	8.272	9.165	0.074

[a, b] Number of common trends, where p is the number of variables and r the number of cointegrating relationships, the rank of the matrix Π

Source: Authors' elaboration

Since the λ_{trace} accepted one cointegrating relationship, whereas λ_{max} did not, the Engle and Granger (1987) methodology was also applied. Results clearly indicate the two prices are cointegrated[8]. As discussed in Section 3, non-linear asymmetries have been proved to be quite common behaviour for a large number of agricultural markets nowadays, including dairy products. Accordingly, TAR and M-TAR models were considered using the residuals from the estimated long-run equilibrium relationship: $\mathrm{P_s} = 0.39 + 0.53\mathrm{P_p} + 0.03\mathrm{D_{s,t}} + \hat{\mu}_{\mathrm{t}}$, where $\mathrm{D_{s,t}}$ represents the shift dummy with $\mathrm{t} = \text{March } 2005 - \text{May } 2012$. Both the AIC and SBC (Schwarz-Bayesian Criterion) indicated the need to include seven lags for $\Delta\hat{\mu}_{t-1}$ in order to exclude any autocorrelation from residuals (Table 4).

Although point estimates for ρ_1 and ρ_2 indicate convergence, i.e., they are all negative and satisfy the condition $(1+\rho_1)(1+\rho_2) < 1$, the Φ and Φ^* coefficients did not reject the null hypothesis of no cointegration. Therefore, neither threshold cointegration nor asymmetric mechanism were detected regarding the organic fluid milk supply chain (see fourth and fifth row in Table 4, respectively).

[8] Engle-Granger test did not accept the null of non-stationary residuals, with a p-value of 0.006 (Tau-statistic: -3.564).

Table 4 - Estimates of Threshold Cointegration for the Organic Milk System

	TAR[a]	TAR consistent[a(*)]	M-TAR[a]	M-TAR consistent[a(*)]
$\boldsymbol{\rho_1}$	-0.111	-0.0742	-0.141	-0.354
	(0.059)	(0.062)	(0.061)	(0.107)
$\boldsymbol{\rho_2}$	-0.159	-0.193	-0.117	-0.105
	(0.069)	(0.065)	(0.067)	(0.049)
$\boldsymbol{\tau}$	0	0.013	0	0.003
$\boldsymbol{\Phi_u}$ and $\boldsymbol{\Phi_u^*}$ $(\boldsymbol{\rho_1} = \boldsymbol{\rho_2} = \boldsymbol{0})$ [c]	3.679	4.628	3.541	6.366
F-test $(\boldsymbol{\rho_1} = \boldsymbol{\rho_2})$ [d]	0.352	1.273	0.088	5.471
			C.V.[b]	
$\boldsymbol{\Phi_u}$ and $\boldsymbol{\Phi_u^*}$	5.519	7.029	6.225	8.567
F-test	3.374	6.909	3.682	8.408

Values between brackets are the standard errors
[a] AIC selected seven lags for the $\boldsymbol{\Delta\hat{\mu}_{t-i}}$ $(\boldsymbol{p_{\max} = 10})$
[(*)] For the two models, the threshold value was estimated through Chan's methodology
[b] Critical Values were simulated for 5% sig. level (1,000 Monte Carlo simulations)
[c] Test for threshold cointegration
[d] Test for asymmetric price adjustment

Source: Authors' elaboration

In order to ensure that the linear model is superior to the nonlinear one, the Hansen (1999) Linearity Test was employed (Table 5). Since the test failed to reject the null hypothesis of linearity, a linear VECM(6) was estimated. Misspecification tests were performed (see Doornik and Hendry, 1997), verifying the statistical adequacy of the model[9].

Table 5 – Linearity Test for the Organic Milk System

	Test	P-Value
1 regime vs. 2 regimes	28.759	0.104

[a] The number of 1,000 bootstrap replications was based on previous work (Ben-Kaabia and Gil, 2007)

Source: Authors' elaboration

Table 6 reports the estimates of the unrestricted and restricted cointegrating vectors, as well as hypothesis tests on long and short-run parameters. The parameter β is usually referred as the long-run price transmission elasticity[10]. However, Lütkepohl and Reimers (1992) and most recently Lloyd et al. (2006, p. 129), argue they are "[…] by construction partial derivatives predicated on the ceteris paribus assumption" and when richer dynamics occur (i.e. a feedback-system in terms of causality) such inference may be not of interest. Moreover, Kinnucan and Zhang's (2015) work heavily contributes to the idea that a price transmission elasticity equal to one does not mean perfect transmission, being this statement inconsistent with the model of Gardner (1975). These authors showed that if the

[9] Results obtained from Breusch-Godfrey for Autocorrelation, Multi-ARCH LM, and Jarque-Bera Normality test were quite satisfactory (For both the Breusch-Godfrey and the Multi-ARCH LM tests we include twelve lags. Results are available upon request).

[10] See, among others: Abdelradi and Serra, 2015; Ben-Kaabia and Gil, 2007, 2008; Brümmer et al., 2009; Busse et al., 2012; Conforti, 2004; Goychuk and Meyers, 2014; Hassouneh et al., 2015; Listorti and Esposti, 2012b; Simioni et al., 2013; Tekgüç, 2013; Verreth et al., 2015.

price changes are caused by shocks on the supply side (i.e., the industrial processor in our specific case), the PT elasticity (EPT hereafter) has to be less than one. Indeed, as clearly proved by Kinnucan and Zhang (2015), for EPT=1 to hold, the retail-demand and the processor-supply curves must have the same elasticity, which would be quite unusual and against the literature's findings on agricultural supply elasticity. The first row of Table 6 shows the unrestricted cointegrating vector, characterized by a β estimate of 51%. As shown in the third row, the test[11] rejects the null hypothesis of long-run price homogeneity [$\beta(1,-1)$] (Table 6).

Table 6 – Organic Milk System: Results from Cointegration Analysis, Restrictions on deterministic, Long and Short-run Coefficients.

Unrestricted cointegrating vector	$\boldsymbol{P_s - 0.51\, P_{\mathrm{p}} - 0.40\mu_1' - 0.04 D_{s,t} = Z_t}$	
Linear Restrictions	**LR-Statistic**	**P-Value**
Long-run Homogeneity $\boldsymbol{\beta(1,-1)}$	5.45	0.020
Weak Exogeneity $\boldsymbol{\alpha_{Pp} = 0}$	0.131	0.720
Exclusion Test[a] $\boldsymbol{\mu_1' = 0}$	4.520	0.033
Exclusion Test[b] $\boldsymbol{D_{s,t} = 0}$	4.880	0.027

[a,b] The last two rows refer to the exclusion tests, which show whether the deterministic terms, i.e., the constant and the dummy, enter the cointegration space or not. The model is well-specified since for neither parameter does the LR statistic accept the null hypothesis of exclusion.

Source: Authors' elaboration

Regarding short-run parameters, weak-exogeneity test on the load matrix indicates the P_{p} is weakly-exogenous (see fourth row), and a cost-push mechanism well describes the dynamic of price transmission. Since the null hypothesis could not be rejected, the processor price does not adjust to deviations in the long-run, and all the adjustment falls on the retail price. The coefficient $\alpha_{Ps} = -0.05$ features a quite slow error correction, i.e., 5% each period, taking around twenty months to resettle the steady-state.

When considering the conventional fluid milk supply chain, the same methodological steps as above were adopted. The graphs suggest the presence of a shift in the series, around March 2010, and four impulse dummies accounting for price spikes[12]. Accordingly, a VAR(4) was specified[13], on the basis of the AIC. Results from the Johansen test indicated that we failed to reject the null of no-cointegration. Hence, as suggested by Juselius (2006), the sample was split into two sub-samples, since structural break may invalidate the assumption of constant parameters. Given the date of the structural change, the two new specimens were defined for the periods January 2001-February 2010 and March 2010-

[11] After normalizing on consumer price, the process is (just-)identified, and further (non-identifying) restrictions on short and long-run parameters are imposed to enable economic interpretability of the results. Because restrictions are asymptotical $\chi^2(v)$ distributed, v being the number of imposed restrictions, the LR statistic is adequate for testing.

[12] The four impulse dummies were $D_{i,m} = 1, m = 1,2,3,4$ at 2006:01, 2008:04, 2008:12, and 2010:03 respectively, and zero otherwise.

[13] For model specification, see (16).

February 2015. Regarding the first sub-sample, a VAR(2) was specified according to the AIC, and a constant was restricted to the cointegrating vector as

$$\Delta P_t = \alpha(\beta' P_{t-1} + \mu_1') + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \varepsilon_t. \qquad (17)$$

Neither λ_{trace} or λ_{max} indicated evidence of cointegration and, accordingly, a VAR in first difference was specified. For the second sub-sample, a VAR(3) was specified based on the AIC, and a constant restricted to the cointegrating space[14]. Results from Johansen's λ_{trace} and λ_{max}, indicated the existence of one cointegration relationship (Table 7), which has the following expression: $P_s = 0.28 + 0.67P_p + \hat{\mu}_t$.

Table 7 – Conventional Milk System: Johansen Co-integration test

$p-r$ [a]	r [b]	λ_{trace}	C.V.	p-Value	λ_{max}	C.V.	p-Value
2	0	39.686	20.164	0.00	18.217	15.892	0.021
1	1	6.366	9.142	0.249	4.362	9.165	0.361

[a,b] Number of common trends, where p is the number of variables and r the number of cointegrating relationships, the rank of the matrix Π

Source: Authors' elaboration

As in the organic milk case, in the second sub-sample, we have tested for threshold cointegration and asymmetries in the price transmission mechanisms using the residuals from the above cointegration vector. Estimated parameters of the Threshold models (TAR and MTAR) are shown in Table 8. The AIC and SBC indicate four lags for $\Delta\hat{\mu}_{t-1}$.

Table 8 – Estimates of Threshold Cointegration for the Conventional Milk System

	TAR[a]	TAR consistent[a]	M-TAR[a]	M-TAR consistent[a]
$\boldsymbol{\rho_1}$	-0.179	-0.122	-0.153	-0.155
	(0.104)	(0.107)	(0.096)	(0.086)
$\boldsymbol{\rho_2}$	-0.230	-0.275	-0.299	-0.419
	(0.106)	(0.100)	(0.129)	(0.179)
$\boldsymbol{\tau}$	0	0.012	0	-0.002
$\boldsymbol{\Phi_u}$ and $\boldsymbol{\Phi_u^*}$ $(\boldsymbol{\rho_1 = \rho_2 = 0})$ [c]	3.402	4.047	3.804	4.333
F-test $(\boldsymbol{\rho_1 = \rho_2})$ [d]	0.138	1.273	0.846	1.778
			C.V. [b]	
$\boldsymbol{\Phi_u}$ and $\boldsymbol{\Phi_u^*}$	5.570	6.817	5.924	8.214
F-test	3.234	6.384	3.577	8.408

Values between brackets are the standard errors
[a] AIC selected four lags for the $\boldsymbol{\Delta\hat{\mu}_{t-i}}$ in (11) $(\boldsymbol{p_{\max} = 10})$
[b] Critical Values were simulated for 5% sig. level (1,000 Monte Carlo simulations)
[c] Test for threshold cointegration
[d] Test for asymmetric price adjustment

Source: Authors' elaboration

The null hypothesis of no threshold cointegration could not be rejected, so the conventional market does not account for asymmetric adjustments as it was the case for the organic market (see fifth and sixth rows of Table 8). Hansen Linearity Tests are again applied to

[14] For the model specification, see (17).

check the superiority of the linear model, and results indicate that the linear model is clearly preferred (Table 9).

Table 9 – Linearity Test for the Conventional Milk System

Hansen Linearity Test	Test	P-Value [a]
1 regime vs. 2 regimes	11.078	0.654

[a] The number of 1,000 bootstrap replications was based on previous work (Ben-Kaabia and Gil, 2007)

Source: Authors' elaboration

Accordingly, a VECM(2) was estimated and misspecification tests performed, which gave satisfactory results (Results are available upon request). Table 10 reports the estimates of the unrestricted and restricted cointegrating vector, as well as hypothesis tests on long and short-run parameters and the retail mark-up.

Table 10 – Conventional Milk System: Cointegration Analysis, Restrictions on Long and Short-run Coefficients and Retail Margin

Unrestricted cointegrating vector	$\boldsymbol{P_s - 0.72\, P_{\mathrm{p}} - 0.28 = \mu_t}$	
Linear Restrictions	LR-Statistic	p-Value
$\boldsymbol{\beta(1,-1)}$	3.49	0.062
$\boldsymbol{\alpha_{Pp} = 0}$	0.04	0.85
Restricted cointegrating vector	$P_s - P_{\mathrm{p}} - 0.24 = \mu_t$	
Retail Margin	$(e^{0.24} - 1) \times \ln P_{\mathrm{p}} \times 100 = 27{,}5\%$	

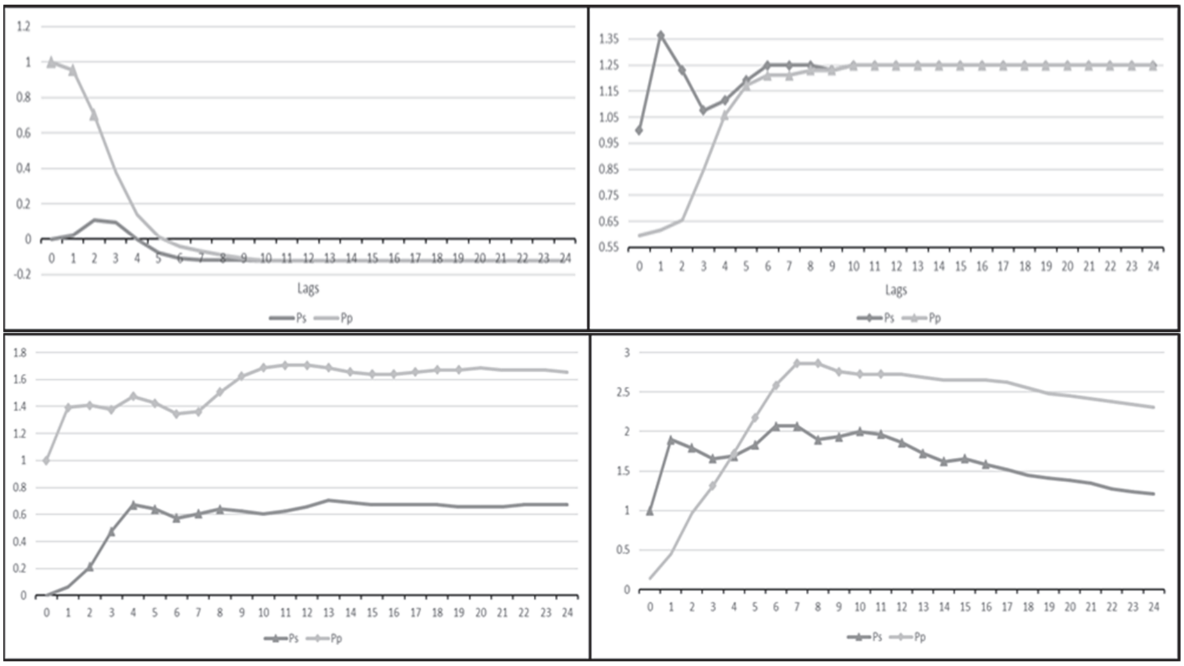
Source: Authors' elaboration

The null hypothesis of long-run price homogeneity cannot be rejected, as reported in the third row of Table 10. Every percentage change in the processor (retailer) price generates a percentage change of the same size and direction in the consumer (processor) price. Additionally, only the retail price adjusts to deviations in the long-run, since the processor price proves weakly-exogeneity (see the fourth row of Table 10). The fifth row reports the cointegrating vector restricted on the long-run price homogeneity. The constant did not vary significantly, whereas the α_{Ps} moved from -0.38, in the unrestricted case, to -0.20 in the restricted one. Hence, the speed of adjustment decreased by nearly 50%, and retail price adjusts 20% of the disequilibrium in each period. The retailer's mark-up on the processor is given by the equation of Tiffin and Dawson (2000), as reported in the sixth row. Given that the average percentage margin for the second sub-sample approaches 24%, such results are consistent. However, no data are available concerning the cost structure of retailers so that no further comments can be made on such results. Besides, this would go beyond the scope of this work.

1.5.2. Short-Run Dynamics

Impulse Response Functions were computed to gain more insights on short-run dynamics for both markets investigated. These show price responses to an unexpected shock to the system. When the VECM(p) is stable, responses are represented as a Moving Average (MA) process, where the MA coefficient matrices contain the IRFs. Nevertheless, the residuals' variance and covariance matrixes are non-diagonal, invalidating the assumption that shocks occur in just one variable, which may generate a misleading picture of existing dynamic relations. In other words, there exists contemporaneous correlation within the system. For this reason, orthogonalized IRFs (i.e., making the error terms uncorrelated) are preferred; they are obtained through the Wald decomposition of the MA representation (Lutkepohl, 2005). Orthogonal IRFs have been then normalized on the estimate at time 'zero' of the impulse variable, and the ordinates' axis displays the unit change (Figure 3). IRFs represent a good indicator of price dynamics along the supply chain.

Figure 3 – Conventional (Upper-panels) and Organic (Lower-panels) System: Orthogonalized IRFs for shocks to both Pp (Left panel) and Ps (Right panel)



*A mark shows when the response is significant
Source: Authors' elaboration

With regards to the organic system, the own-price response to shock on the processor side is positive and significant as expected, and it persists significantly up to the 19th period. Retail price positively responds after two months and lasting until the 8th month. The system then reaches a new equilibrium, and all prices respond significantly to any shocks. Own response to shocks on the consumer side, is, again as expected, positive and significant until the 16th period. The response of processor price is positive and significant starting from the third month, overcoming the initial shock and reaching its maximum after seven periods.

After reaching the apogee, it starts decreasing and lasts significantly until the 11th month. Finally, since the magnitude of the response is greater when a shock occurs on the demand-side, a demand-pull mechanism seems to explain price transmission behavior in the organic sector better.

When considering conventional milk, especially for the first sub-sample, all the computed IRFs were non-significant, and prices move independently of one another[15] (according to IRFs results, the system was not cointegrated). However, regarding the second sub-sample, when a shock occurs on the offer side, both prices return to the steady-state after eight months approximately. However, only the response of processor is significant and positive, and only for the first quarter, after which it starts decreasing and becomes negative, co-moving again with the retailer price. The retailer price, however, does not show any statistically significant response, although it displays a very moderate and positive response for the first four months and later becomes negative. When the shock is on the demand side, the own price response is quick and positive and reaches the equilibrium after about eight months. Likewise, Pp reacts positively but becomes significant only after four months.

1.6. Discussion

Comparison between the two fluid markets reveals some significant differences. Firstly, for the organic supply chain, results showed a more responsive market than the conventional one, since cointegration has been found for the whole period. In the case of the conventional fluid milk market, prices only seem to co-move since 2010. Before that date, results suggest that prices at the retail (processor) level were mainly determined by the retail (processor) supply and demand and not by the price paid to processors (sold at the retail level). Therefore any unexpected shock occurring on any side of the chain was not transmitted either upstream or downstream. After 2010, processor and retail process co-move. The comparison of results from each market indicates that the elasticity transmission and the speed of adjustment parameters are higher in the conventional market (the speed of adjustments after an unexpected shock in the organic market is four times lower than in the conventional counterpart). Probably this has to do with the different market structure of the organic and the conventional fluid milk. While supermarkets and hypermarkets concentrate a significant share of the final demand for conventional fluid milk, in the case of the organic product alternative marketing channel play a much relevant role, reducing the market share of supermarkets and hypermarkets. Borrowing the hypothesis number three in Borenstein et al. (1997), we can argue that search costs for the organic milk consumer are higher, in the way that when prices change, the expected gain from searching other distributors is smaller than for their conventional counterpart. Therefore, this may reduce (temporarily) the

[15] Because of space limitations, IRFs graphics are not presented here, but are available on request.

elasticity of the consumer demand and dampen the speed of adjustment. The yet limited awareness of the consumer to organic food, in general terms, together with a still restricted access to information for this niche market could explain such higher search cost. Moreover, because informal distributors (i.e., farmer-markets and direct sales) represent a significant part of organic-food retailing, the consumer may have to engage an intense research to understand if price changes are due to market conditions or simply marketing strategy.

Short-run dynamics show that for the organic system, an increase in the consumer price causes a larger increment on the processor side; this suggests supply is inelastic (as expected, an increase on the demand side produces a larger increase on the production side, since the quantity produced is fixed in the short-run), and upstream chains are better organized. Organic milk producers are fewer in numbers, improving their bargaining power and allowing for a better management on the supply side, improving the capability of coping with adverse market situations. Moreover, the processor reacts quicker than in the conventional case, and the magnitude of the response is significantly higher. When shocks take place on the supply-side, the consumer price response is significant, and a new equilibrium is reached. Changes in regulations, production methods or quality standards affect the cost-structure of producers, and shock may become permanent.

In the conventional fluid milk market, after 2010, we observe a certain stickiness in retail price, which may reflect the fact that when shocks occur on the supply-side, the retailer absorbs the shock without transmitting it to the customer. They then probably exert market power to pressure the processor to lower their selling price, which explains the rapid decrease of processor response. When demand-side shocks occur, the retailer tends to increase the volume purchased, causing an increase in processor price and a quick decrease in consumer price. Since supply shortages of fluid milk are exceptional, we consider here only positive shocks due to an increase in demand. The delay in processor response may be a consequence of private agreements; for instance, only when the milk spot price exceeds a certain threshold, previously agreed by the two actors, does the price paid to processor increase, since the processor cost structure changes substantially[16]. The most recent studies on the fluid milk sector many often conclude the price transmission dynamics is asymmetric (see Table 1). Nevertheless, our results are consistent with those of Serra and Goodwin (2003) for the Spanish market, whose characteristics are quite similar to those of the Italian sector, and with those of Bakucs et al. (2012) for Hungary. Being the fluid milk a low value-added product, this could prevent APT and tighten relationships between industrial processors and retailers (Serra and Goodwin, 2003). However, as indicated in Table 1, this comparison has to be cautious: each study is country-specific, and

[16] The SPOT milk price is the reference price for crude-milk at the farm-gate in Italy, i.e. the reference price for farmers.

the vertical relationship investigated varies across studies, as well as the direction of causality. Looking at the speed of adjustment (see the last column in Table 1), they are distributed within a quite wide band, from 2% to 39% of shock transmitted in one period. One might say they are strongly dependent on the market structure and agents characteristics. Accounting for this, the two speed of adjustment we found for the conventional and the organic milk sector in Italy, are within the band (20% and 5% respectively), despite organic milk prices are not considered yet in any study, up to the authors' knowledge. According to Serra and Goodwin (2003), Peltzman (2000), and Santeramo and von Cramon-Taubadel (2016), high-perishable products reveals a (more) symmetric PT mechanism. Kim and Ward (2013), and Ward (1982), despite they both found APT in perishable products, they describe how retailers may be reluctant to increase prices fearing sales reduction and, hence, increase spoilage.

1.7. Conclusions

The main objective and contribution) of this paper was to analyze price dynamics in a quality-differenced market, in relation to its conventional counterpart. The organic fluid milk market has been chosen as the case study. The methodological approach has been based on time series econometrics, with a special focus on asymmetric price dynamics.

Results suggest a number of points. First, in the case of the Italian fluid milk market (either organic or conventional), price reactions to unanticipated supply and demand shocks are symmetric, mainly for two reasons: i) the monthly frequency of our dataset, which contributes to smooth the price series; and ii) in spite of the high degree of perishability of milk, most of the vertical relationships between farmers and processors are based on long-term contracts. Second, in both markets, all the adjustment process takes place on the consumer price, a result consistent with the literature. In fact, (see also Table 1, fifth column) the lion share of PT-related works found causality running from upstream levels to retailers, and not vice versa. Therefore, only retail prices adjust when upstream prices change, while the latter are often exogenous so that they do not adjust to changes in consumer prices. This finding is consistent with the general rule stating that agricultural supply is inelastic in the short term. However, the magnitude and the speed of the adjustment to a new equilibrium differ in both markets as both are characterized by different market structures. In the conventional market, the higher retailer-concentration may oblige distributors to readjust quickly to the equilibrium, since it would be easier for the consumer to switch from one retailer to another. In the organic market, the supermarkets' market share is less significant, and alternative marketing channels are available, generating a slower adjustment since heterogeneous distributors prevent a quick pass-through.

Short-run dynamics depict a more favorable situation for organic milk, and both the magnitude and the speed of the responses overcome those on the conventional side.

Asymmetries do not appear to exist between the last two levels of the supply chain. However, asymmetries may exist between the farm-gate and the other economic agents along the supply chains, which should be explored in the future when reliable data become available. Moreover, results should be interpreted with caution as this study was based on scanned data from a single retailer and generalization could be problematic. In any case, this study opens new opportunities for further research as new studies about price dynamics in quality-differenced products are necessary to understand better how different market and governance structures influence the performance of their respective supply chains.

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2. Vertical Price Transmission in Milk Supply Chain: Market Changes and Asymmetric Dynamics

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Abstract

The milk supply chain undergone great changes during the last two decades and, in particular, there has been a great trade liberalization and a conspicuous policy change due to the abolition of milk quotas. These market changes are likely to have had a large impact on the transmission of price within the supply chain. Our analysis aims at exploring how market changes have altered vertical price transmission, and, in particular, how asymmetries have changed over time. We use an asymmetric error correction model to infer on short-run and long-run adjustments and conclude on the potential role played by market changes.

Keywords: Asymmetries, Error Correction Model, Price transmission, Supply Chain

2.1. Introduction

The Italian dairy industry, especially the fluid milk sector, has been under the spotlight for the last ten years, with farmers' unions leading numerous public demonstrations claiming an unfair value distribution along the chain and more often to be producing at a loss. Moreover, the abolishment of the quota regime enhanced their discontent, since they claim the (potential) increase in production will plunge the price, and competitors with largely lighter cost structures may enter the market. The European Commission (EC) stated that the abolishment of the quota regime is necessary for the enhancement of the farm-gate competitiveness and a more market-driven dairy industry, representing a very significant step toward the liberalization of the milk sector. Indeed, starting in 1992 with the McSharry Reform, the Common Agricultural Policy (CAP hereafter) witnessed several interventions in the last decades aimed at a dramatical reduction of the EU's protectionism over its agricultural markets. Concerning the milk sector, the analysis of both the economic calendar and market fundamentals point to the year 2007 as the watershed in the European dairy industry, and, consequently, on the Italian domestic market.

The analysis of vertical price transmission (VPT) along food chains has attracted considerable interest among agricultural economists (see Kouyaté and von Cramon-Taubadel (2016) for a thorough analysis of the related literature). Indeed, the price is the primal link between economic agents, representing a good measure of supply chain integration and, thus, its efficiency (Abdulai, 2002; Goodwin and Holt, 1999; Lloyd, 2016; Serra and Goodwin, 2003). However, despite the broad application of VPT analysis on several food supply chains, a small number of applications so far investigated the impact of CAP reforms on price transmission dynamics, especially concerning the dairy sector.

We tried to fill the gap in price transmission (PT) literature exploring its dynamics along the Italian fluid milk supply chain. Two price series are employed, featuring the price paid to industrial producers by retailers and the price paid by consumers for fresh fluid milk, covering 16 years, from 2000 to 2016. The applied methodology permits discovering the degree of the PT mechanism, the speed at which a shock is pass-through, the nature of transmission (i.e., cost-push, demand-pull, or feedback system), and whether the transmission is (a)symmetric. Since non-structural time series models are applied, a thorough understanding of both the market structure and the political events characterizing the analyzed period is required. Therefore, a review of international, European and domestic measures deployed within the 2000-2016 period is provided in Section 2, along with a brief literature review of the most recent studies on PT in agrifood markets in Section 3. Section 4 details the applied methodology, while the nature of the data collected, together with results, are presented in Section 5. Finally, Section 6 concludes.

2.2. The Italian Dairy Sector and the EU Policy

The whole fresh fluid milk market is one of the most important agri-foods in Italy. The dairy industry is a major asset which is worth 15 billion euros, 11.4% of the food industry turnover, with more than two thousand dairy firms involved and a workforce of thirty thousand (Gonano and Mambriani, 2016). Moreover, Italy is the 6th EU-28 fluid (cow) milk producer, providing 7% of the total EU production (EUROSTAT, 2017).

However, Italy is a net importer of dairies (in milk equivalent), and as illustrated in Table 11, fluid milk is the mostly imported category.

Table 11 – Italian Dairy Trade Volumes (in milk equivalent), 2015

Product	Import	Export	Balance
Bulk Milk	2.015.172	43.700	- 1.971.471
- Packaged	*460.896*	*24.430*	*- 436.466*
Cheeses	511.082	363.158	- 147.924
Milk Powder	103.968	12.987	- 90.980
Butter and Milk's Fats	71.150	9.435	- 61.716
Whey	71.328	399.013	327.685

Source: Authors' elaboration on ISTAT.

The Italian dairy sector is oriented to cheese production. According to last figures, nearly 70% of the total produced milk is destined to cheese production, of which approximately 50% labeled as GI, and only 16% intended to fluid milk production (Gonano and Mambriani, 2016). Therefore, the dairy sector is of central importance when referring to the Italian agribusiness, and the recent uprising all over the Italian territory unveiled an increasing discontent within the industry, with the CAP heavily criticized for its (new) liberal policy.

2.2.1. EU and International Policy Measures: the Liberalization of the European Dairy Market

In 1994, at the end of the so-called Uruguay Round, the Marrakech Protocol was signed, and GATT was formally taken over by the World Trade Organization (WTO) on 1st January 1995, together with the Agreement on Agriculture (AoA). What WTO intended was to initiate the liberalization of the world agricultural markets (Swinbank, 2016), and certainly, this has shaped the reform process the CAP has been going through, particularly from 1992 up to 2008 (Cunha and Swinbank, 2011; OECD, 2011; Swinbank, 2016). Heavy-regulated European agricultural markets - the dairy, cereals, fruit&vegetables, and sugar sectors - witnessed a progressive softening – elimination in some cases - of all those market measures classified as distortive for the international trade (i.e., export subsidies, price support, reference/intervention prices, import tariffs, and aid coupled to production) (Assefa et al., 2016; Bouamra-Mechemache et al., 2009; Gohin and Latruffe, 2006; Kloosterboer, 2016; Meijerink and Achterbosch, 2013; Swinnen et al., 2014).

Since our analysis focuses on the dairy sector, is with the reform of 2003 (usually referred as both the Fischler or Luxembourg Reform) the milk market experienced the major shift towards a free market (Bonnet et al., 2015; Gohin and Latruffe, 2006). As stated by the European Court of Auditors (2009, p. 13) "[…] In accordance with the decisions contained in the Agenda 2000 action program, the 2003 reform initiated the liberalization of the milk sector by reducing price support and creating direct income support". Indeed, the milk target price was abolished, intervention prices for dairy products lowered, and national milk quotas increased by 1.5%. In more details, the reform set a 25% decrease in the butter intervention price over four years, from 2004/2005 to 2007/2008, and a further 15% reduction for the SMP over a three-year period, from 2004/2005 to 2006/2007 (Bouamra-Mechemache et al., 2009; Kloosterboer, 2016). Also, the dairy premiums and other coupled additional payments have been included in the (new) Single Payment Scheme (SPS) from 2007, and the decision of dismantling the milk quota system was taken (DG-AGRI, 2010; Gohin and Latruffe, 2006)[17]. Likewise, the new Single CMO (Common Market Organization)[18] signed in 2007 has had a significant impact on the European milk market, amending the first CMO for dairy products established in 1999 (Council Regulation (EC) No 1255/1999). New intervention prices for butter and skimmed milk powder (SMP) were introduced:

[17] As pointed in Gohin and Latruffe (2006), coupled direct payments were blue box measures under the WTO's, and moving to a more decoupled aid's framework means green box measures.
[18] Council Regulation (EC) No 1234/2007.

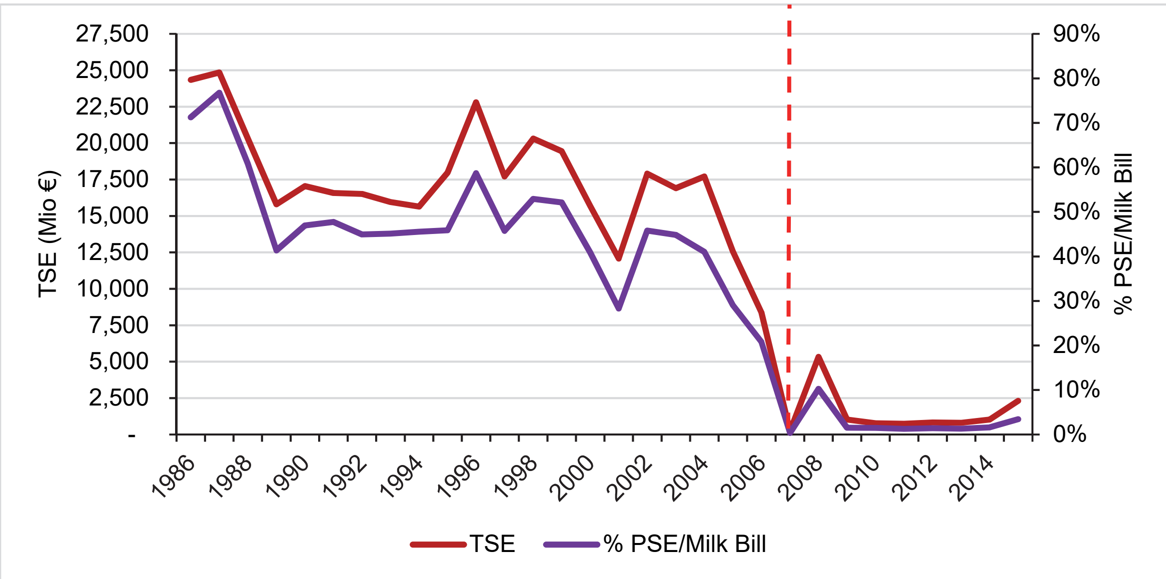
221.75[19]€/100 Kg from the 2007/2008 campaign and 174.69€/100 Kg from the 2006/2007 marketing year, respectively[20] (See Figure 6). Furthermore, the threshold quantity above which the buying-in for butter has to be carried out by a tender procedure has been lowered by 10,000 thousand tons, i.e., 30,000 tons in 2007. Moreover, for the period June 2007-August 2011, export subsidies for milk were set to zero due to high world prices (except for the period January-November 2009) (DG-AGRI, 2011; Meijerink and Achterbosch, 2013). One of the most important elements of the European dairy policy was the quota system. Introduced in 1984[21] to regulate the structural surplus generated by the system of guaranteed price and reducing the pressure of the milk sector on the CAP's budget, they were dismantled on March 1st of 2015 (Giles, 2015; Tonini and Domínguez, 2009). Milk quota, by definition, permitted to maintain a sort of price support for dairy farmers[22], generating a positive spread between the European and World milk prices. Eventually, is with the implementation of the Luxembourg Reform (2003) first, and the Health Check (2008)[23] afterward when the EU ultimately decided the dismantling of the quota system by March 1st, 2015, being the 2014/2015 the last quota-milk campaign (Anania and Pupo D'Andrea, 2015). However, a "soft-landing" phasing-out measure was adopted in the 2008 reform, aimed at preventing the dramatic drop in milk price due to a potential increase in production. Therefore, quota-regimes were augmented from the 2008/2009 campaign by 2% and 1% from the 2009/2010 campaign for five consecutive years, except for Italy for which the cumulative 5% increase was frontloaded on the beginning of the 2009/2010 campaign[24].

Relative to the dairy sector only, the Total Support Estimate (TSE)[25] represented in Figure 4 is an OECD indicator that combines all kind of agriculture-related public monetary expenses. This decreased from 2.63% of GDP in the 1986-88 period to 0.84% in 2007-09. Furthermore, the share of the PSE (Producer Support Estimate, the direct support given to agricultural producers) over the gross farm milk bill witnessed a consistent decrease in 2007, reduced almost to zero and maintaining a meager impact afterward. The two illustrated indicators cement the idea that CAP has moved to a far less-intervening policy framework, consistent with the liberalization course that invested agricultural markets, and

[19] This is the corresponding 90% of the reference price, set at 246.39€/100 Kg.
[20] The threshold price for SMP was lowered to 169.8 €/100 Kg since September, 2008.
[21] See the two Regulations 856/84 and 857/84
[22] The so-called quota-rent: the amount of rent generated from a restriction on supply (Tonini and Domínguez, 2009).
[23] The European Commission released in November 2007 the document "Communication from the Commission to the European Parliament and the Council. Preparing for the Health Check of the CAP Reform".
[24] See the Report from The Commission to The European Parliament and The Council: Evolution of the market situation and the consequent conditions for smoothly phasing-out the milk quota system - second "soft landing" report (COM(2012) 741 final).
[25] See the OECD's PSE Manual at http://www.oecd.org/tad/agricultural-policies/psemanual.htm.

particularly the dairy sector, from 1992 with the McSharry Reform. However, the year 2007 seems to fairly represent the moment in which the European dairy industry experienced a major and abrupt change toward the free market. Indeed, most of the instruments introduced with the Fischler Reform became active in 2007, together with the CMO Reform.

Figure 4 – Total Support Estimates (TSE) for Milk in the EU and its Composition, 2000-2015



Source: Authors elaboration on OECD data

2.2.2. The Italian Dairy Sector: Facts and Figures

We described so far the CAP reforms and which intervention have affected the dairy sector. In this section, a brief analysis of the Italian market is provided, describing the effects of the mentioned policies on the Italian domestic market.

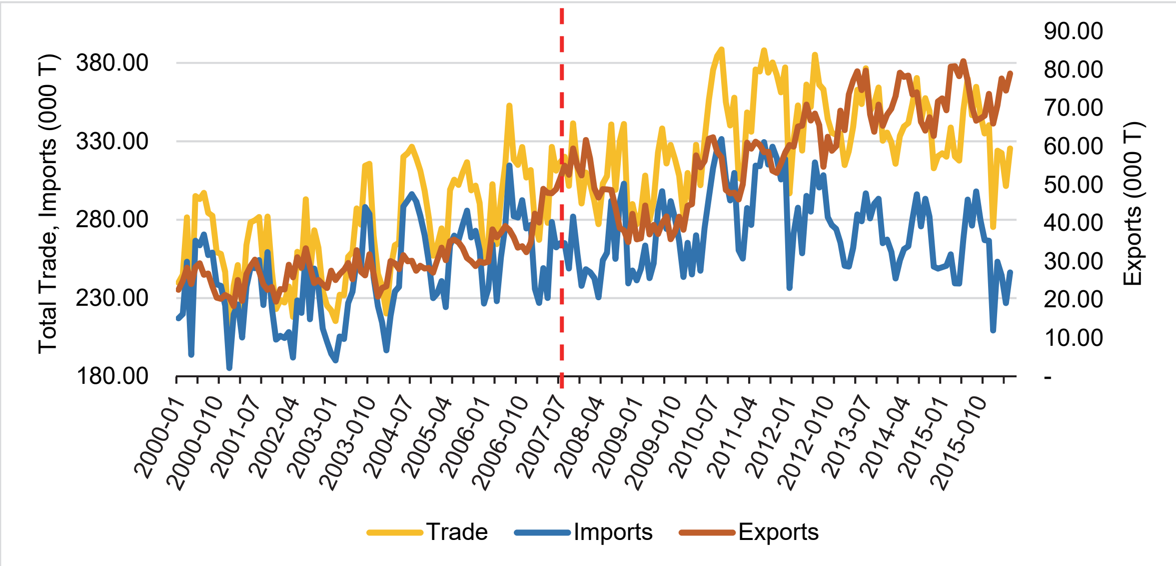
Figure 5 illustrates imports and exports trends for dairy products in Italy. Considering the year 2007 as the breaking point, we analyzed the behavior of Italian trade volumes before and after. While imports seem to maintain the same pattern with a very tenuous positive growth of about 11%, exports entail a more complex course. Indeed, the price surge in 2007 together with an increasing demand for protein in emerging markets (i.e., China and India mostly) may explain the massive increase in exports (+110%), which is ten times bigger than that of imports. Table 12 illustrates the average trade flows for the two selected periods regarding five macro-categories of dairy products, confirming the structural change occurred for the Italian trade flow. Focussing just on the fluid milk category, while imports do not experience a significant change in imports (+2.5%), exports did, doubling that of the pre-2007 scenario. The trade volume for fluid milk increased solely by 4.5% since imports have a greater weight than exports do. However, the overall trade in dairy products rose by 21,5 %.

Table 12 – Average Trade Flows in Italy, 2000-2016 (000 Tonnes)

	Period	Milk and Milk Cream	Yogurt and Other	Whey	Butter	Cheese	Total
Exports	2000-2006	15.5	3.8	102	13.3	206.1	68.1
	2007-2016	53.5	5.6	348.3	11.1	299.9	143.7
	% Change	245.90%	46.60%	241.60%	-16.20%	45.50%	**110.90%**
Imports	2000-2006	2,305.86	131.16	53.41	50.89	380.27	584.32
	2007-2016	2,367.50	213.72	121.03	60.81	481.17	648.84
	% Change	2.70%	62.90%	126.60%	19.50%	26.50%	**11.00%**
Total Trade Volume	2000-2006	1,160.66	67.49	77.69	32.09	293.17	3,262.20
	2007-2016	1,210.49	109.66	234.68	35.98	390.52	3,962.64
	% Change	4.30%	62.50%	202.10%	12.10%	33.20%	**21.50%**

Source: Authors' calculations on ISTAT

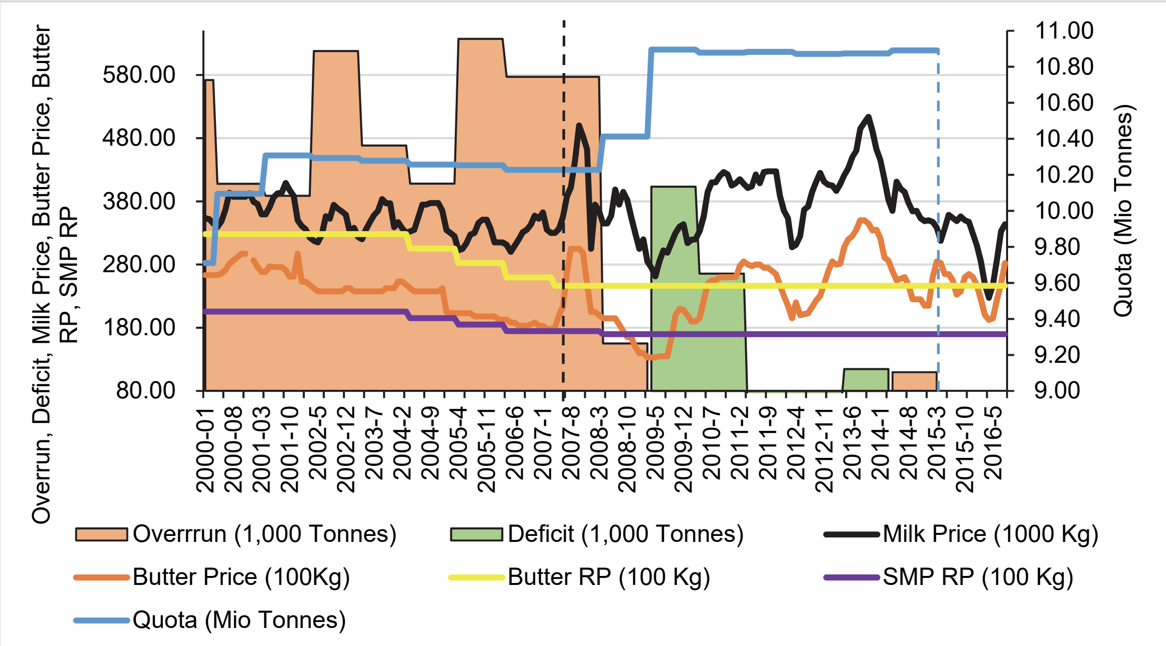
Figure 5 – Italian Trade Volumes, 2000-2016



Source: Authors' elaboration on ISTAT

Figure 6 shows the main policy interventions occurred on the Italian milk market. The blue line indicates the amount of milk quota assigned to Italy (the hatched line define the end of the quota regime), while red (green) bars refer to overruns (deficit) regarding the assigned quota level. It is evident how, particularly from the 2008/2009 Campaign, surpluses reduced dramatically, in reason of a higher quota threshold settled by the Health-Check' soft-landing measure. Indeed, for four consecutive campaigns (i.e., from 2009/2010 to the 2013/2014 marketing year), Italy milk deliveries were under the assigned quota. The two purple and yellow lines represent the two reference prices (RP) for SMP and butter, respectively. Looking at the black hatched line, it is evident how a change occurred around mid-2007, especially when considering the two Italian spot prices for milk and butter (i.e., continuous black and orange lines, respectively). The former lost the cycle component that characterized its course before the break, thus turning far more volatile, as well illustrated by price spikes and valleys on the right-hand-side of the hatched line. Concerning the butter price, it seems quite independent of the raw milk price in the pre-break period, following the general decreasing trend that characterizes its reference price. Nevertheless, after the break, it shows an entirely different pattern, strictly connected with the spot milk price and experiencing many abrupt changes, pointing to a sharp increase in volatility too. Therefore, one might conclude that the liberalization process impacted the Italian domestic market, especially when looking at the spot prices of the two major dairy commodities. Exports were the most affected by the new policy setting, whereas when considering both the imports and the global trade volumes, we spot a little positive growth.

Figure 6 – Italian Spot Milk Price and EU Policy, 2000-2016



Source: Authors' elaboration on MMO (Personal Communication), MMO (Press Release, Various Years within 1999-2016) and the Chambers of Commerce of Lodi.

Table 13 illustrates price volatility[26] characterizing each sub-period for both consumer (P_c) and producer (P_p) price series, setting the breakpoint in August 2007. Results show a substantial increase in volatility for both prices: while consumers experienced an increase of about 28%, producer price's volatility rose by 83%. Margin increase occurs whenever consumer (producer) price increases (decreases), whereas a decrease in margin describes the opposite situation. Given the hatched line representing the structural break, is evident how the pre-break period featured a positive margin change, increasing at a quicker pace when compared to the post-break scenario.

[26] Volatility has been calculated as: $\sigma_t = \sqrt{\frac{1}{m}\sum_{i=1}^{m} r_{t-i}^2}$, where $r_t = 100 \cdot (P_t - P_{t-1})/P_{t-1}$, m is the number of observations and the mean of returns $\bar{r}$ is assumed to be 0.

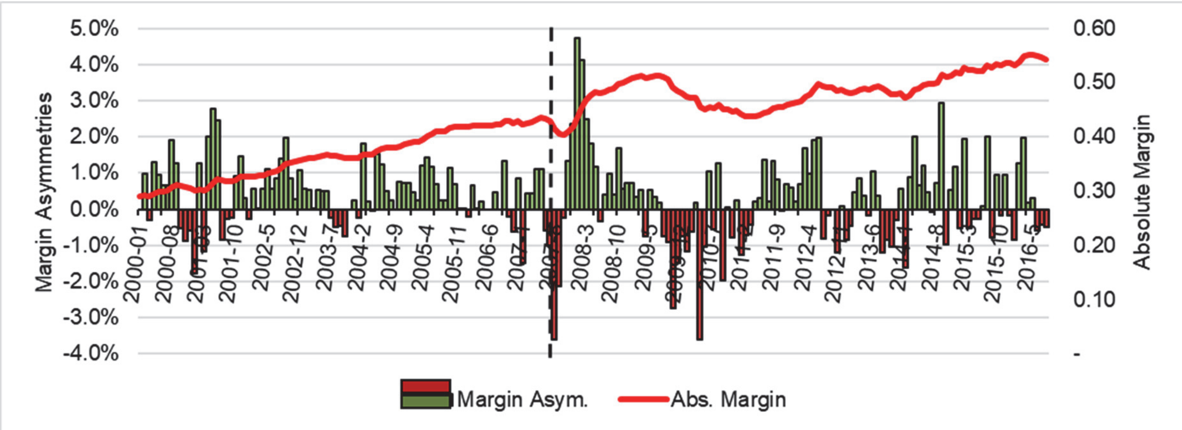
Table 13 – Estimated Volatility, January 2001 – July 2007 and August 2007 – June 2016

	P_c	
Jan 2000 – Jul 2007		0.33
Aug 2007 – Aug 2016		0.42
	P_P	
Jan 2000 – Jul 2007		0.37
Aug 2007 – Aug 2016		0.69

Source: Authors' calculation

Effectively, the margins' average change for the first period doubles that of the second one. Both maximum and minimum values are higher in the second term, probably due to larger volatility in both upstream and downstream prices. However, in absolute terms the average-margin increased in the post-break, being 0.49€/l against 0.37€/l of the pre-break period.

Figure 7 – Italian Milk Margin Trend



Source: Authors' elaboration

Concerning the persistence of margin's changes, looking at the right side of the hatched line there exists a much more equilibrated margin changes dynamics. Green bars, representing margin increases, are much less persistent in terms of months, being the average of positive changes 5.4 months and 3.6 months concerning pre and post-break periods, respectively. On the same line, more than 60% of negative margin changes happened after the break.

2.3. Literature Review

Being prices the instrument by which information is conveyed to all stages of the supply chain, economists are heavily interested in how price shocks are transmitted within (vertical price transmission – VPT) and between (horizontal price transmission – HPT) markets. Gardner' seminal paper on PT (1975) explicates the relevant theory, clarifying '[…] the mechanisms and their implications in an area that has been characterized by its fair share of misunderstanding and overreach' (Lloyd, 2016, p. 1). Since then, a vast literature on price transmission (PT) has been developed and as Kouyaté and von Cramon-Taubadel (2016) pointed out, are 492 the studies embodying this kind of analysis. Focussing on Asymmetric-VPT (AVPT), its importance is related to the capacity of giving a measure of the behavior of economic agents and of its functioning (Ben-Kaabia and Gil, 2007; Lloyd, 2016; Serra and Goodwin, 2003). In recent years, PT analysis witnessed a wide development of new econometric modeling techniques that allow for testing the presence of asymmetries in the PT process (APT). Therefore, new technics together with structural changes occurring in the food industries (i.e., mergers and acquisitions and policy changes), enhanced the interest of economists in founding asymmetric price dynamics. Asymmetries generate a disruption in welfare distribution within the considered supply chain since depending on the sign of the price changes (i.e., positive or negative), the magnitude and the speed of adjustment to the equilibrium may differ (European Commission, 2009; OECD, 2015). In his inspiring study Peltzman (2000) demonstrated that prices rise faster than they fall (i.e., the rockets and feathers phenomenon), spawning APT-related research. Indeed, he found that asymmetries are the rule rather than the exception, highlighting an existing gap in economic theory. More recently, Bakucs et al. (2014) also found APTs in more than 50% of investigated agricultural markets. However, despite a substantial number of works investigate asymmetries, drawing conclusions that should motivate such (inefficient) market behavior and suggest general policy intervention is yet a troublesome task (Vavra and Goodwin, 2005). Nevertheless, Meyer and von Cramon-Taubadel (2004) detailed a thorough description of what may cause an asymmetric price adjustment. Market power is probably the most quoted one (Bailey and Brorsen, 1989; Lloyd et al., 2006; Madau et al., 2016; McCorriston et al., 2001; Sckokai et al., 2013; Sexton, 2013; Shrinivas and Gómez, 2016; Simioni et al., 2013; Soregaroli et al., 2011; Verreth et al., 2015), although not always asymmetries match with high concentrated markets (Acosta and Valdés, 2014; Bakucs et al., 2014; Bettendorf and Verboven, 2000; Peltzman, 2000; Sckokai et al., 2013; Serra and Goodwin, 2003). Ward (1982) was the first scholar linking VAPT with the level of product perishability, concluding that retailers respond more to decreasing wholesale prices than increasing ones. These findings may suggest that retailers are adverse to increase prices for the perishable product since this could lead to sales reduction and increase spoilage. On the other hand, Heien (1980) argues that prices of perishable products are more dynamic, though changing prices is less of a problem. Peltzman (2000) found weaker evidence of asymmetries for perishable products and, alike, Serra and Goodwin (2003) found APT in long shelf-life dairy products while symmetric PT in high-perishable milk products. Kim and Ward (2013) observe that PT in fruit and vegetable commodities is negative asymmetric, and decreases in the wholesale prices are passed through more quickly than increases. Santeramo (2015), in his study on the tomato and cauliflower sectors in Europe, supports both Kim's and Ward (2013) and Ward (1982) conclusions since wholesalers price decreases have a larger impact on retail than price increments. Finally, Santeramo and von Cramon-Taubadel (2016) analyze different products from the fruit and vegetable category and conclude APT is found in 17 out of 40 cases, of which 16 products were classified as “low-perishable,” supporting the hypothesis of a more symmetric PT in high perishable products.

Moreover, there is a vast number of studies investigating how a change in government policies could affect price transmission dynamics. Kinnucan and Forker (1987) studied how government support to producer prices (e.g., floor prices) could cause APT in the US dairy sector. Santeramo and Cioffi (2012) and Cioffi et al. (2011) studied the effects of the EPS (i.e., entry price scheme) in the fruit and vegetable sector in the EU, concluding the stabilization effect on domestic prices is rather limited. Lee and Gómez (2013) estimated how the end of the coffee export quota system (EQS) affects PT between international and retail prices in France, Germany, and the US. They found that retail became more responsive to changes on the international side in the post-EQS period, despite short-run asymmetries and a decrease in the speed of adjustment to the equilibrium. Cacchiarelli et al. (2016) investigated how the mid-term reform of the CAP in 2005 affected the PT process within the wheat-pasta chain. If the farm-wholesaler relationship became symmetric in the post-reform, the opposite occurs in wholesaler-retailer relation, where there is a significant asymmetric long-run behavior from retailers. Han et al. (2016) compared PT behaviors in US cattle markets pre- and post-EPA (i.e., Energy Policy Act), which increased the production of corn ethanol, finding a lower integration and a slower transmission between the investigated markets in the post-EPA period. Esposti and Listorti (2013) determined whether and how temporary trade-policy measure applied to mitigate price bubbles (i.e., the suspension of the European import duties on cereals) do have an impact on PT process in the Italian and North American markets, in particular for cereals. When effective, such policy measure mitigated the impact of the price bubbles. Brümmer et al. (2009) analyzed the PT between wheat and wheat flour in Ukraine during a period of significant policy intervention, and they found a strong coincidence between regimes of high uncertainty and policy interventions, concluding these may amplified instability. Ihle et al. (2012) explored simultaneous impacts of policy reforms and animal health crisis on HPT among four main European markets, showing these significantly impacted PT process in the investigated markets.

Further causes for APTs were explored, such as substitutability between agricultural and other marketing inputs (Bettendorf and Verboven, 2000; McCorriston et al., 1998), adjustment and transportation costs (Azzam, 1999; Chavas and Mehta, 2004; Santeramo, 2015), asymmetric information (Bailey and Brorsen, 1989), and inventory costs (Reagan and Weitzman, 1982).

2.4. Methodology

The first specification for modeling asymmetric price transmission was designed by Wolffram (1971) and later modified by Houck (1977) and Ward (1982), where the response of consumer (retail) price P_c to a shock in processor price P_p was estimated via the model:

$$\Delta P_{c,t} = \gamma_0 + \sum_{j=1}^{m}(\gamma_j^{+}\mathrm{D}_t^{+}\Delta P_{p,t-j+1}) + \sum_{j=1}^{n}(\gamma_j^{-}\mathrm{D}_t^{-}\Delta P_{p,t-j+1}) + \mu_t \quad (1)$$

where $\Delta P = P_t - P_{t-1}$, D_t^{+} and D_t^{-} are dummy variables for positive ($P_{p,t} > P_{p,t-1}$) and negative ($P_{p,t} < P_{p,t-1}$) values, respectively, γ_0 is the constant term and μ_t are error terms. In this context, the hypothesis of symmetric price transmission can be tested against asymmetric adjustment ($\mathrm{H}_0: \gamma^{+} = \gamma^{-}$; $\mathrm{H}_a: \gamma^{+} \neq \gamma^{-}$). However, the model expressed in (1) is not consistent with cointegration (Engle and Granger, 1987) and it neglects the time series properties of the data, such as autocorrelation and unit-root (von Cramon-Taubadel, 1998). The Error Correction Model (ECM) early introduced by the seminal work of Engle and Granger (1987) is linear by definition:

$$\Delta P_{c,t} = \alpha_0 + \beta_0\Delta P_{p,t} + \alpha_1(P_{c,t-1} - \alpha_0 - \beta_1 P_{p,t-1}) + \sum_{i=1}^{k-1}\Gamma_i\,\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\psi_i\,\Delta P_{p,t-i} + \varepsilon_t \quad (2)$$

where $P_{c,t-1} - \alpha_0 - \beta_1 P_{p,t-1} = ECT$ (Error Correction Term), given the cointegrating relationship $P_{c,t} = \alpha_0 + \beta_1 P_{p,t} + u_t$. Hence, all the parameters in (2) are assumed to be constant and whenever a structural break, such a new policy, is considered, such assumption may be misleading. Effectively, one may be interested in a model' structure that considers the behaviour of the price transmission dynamics before and after the structural change to occurr, namely regime-dependent or 'piecewise' linear models (Hassouneh, Von Cramon-Taubadel, et al., 2012). Granger and Lee (1989) introduced the Asymmetric Error Correction Model (AECM), segmenting the Error Correction Term (ECT) into positive and negative values. Generalized by the seminal paper of von Cramon-Taubadel (1998) investigating price transmission dynamics in the German pork market, the AECM is specified as

$$\Delta P_c = \alpha_0 + \beta_0\Delta P_{p,t} + a_1^{+}ECT_{t-1}^{+} + a_2^{-}ECT_{t-1}^{-} + \sum_{i=1}^{k-1}\Gamma_i\,\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\psi_i\,\Delta P_{p,t-i} + \varepsilon_t \quad (3)$$

where ECT_{t-1}=$ECT_{t-1}^{+} + ECT_{t-1}^{-}$, i.e. it is split, into its positive and negative values, and an F-test can be used testing the null of symmetry ($a_1^{+} = a_2^{-}$). Therefore, depending on the sign of the ECT, one may consider (3) as regime-dependent: one regime of linear price response when $ECT_{t-1} > 0$ and another regime of linear response when $ECT_{t-1} < 0$ (i.e., piecewise linear).

Recently, Santeramo and von Cramon-Taubadel (2016), Alam et al. (2016), Acosta et al. (2014), and Capps and Sherwell (2007) also employed the (Vector)AECM for studying APTs within and between different agricultural markets. According to Lee and Gómez

(2013), investigating asymmetries in the short-run requires lagged differenced prices to be split into their positive and negative values, such that:

$$\Delta P_c = \alpha_0 + \alpha_1 ECT^{+}_{t-1} + \alpha_2 ECT^{-}_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_1 \, \Delta^{+} Pp_{t-i} + \sum_{i=1}^{k-1} \psi_2 \, \Delta^{-} Pp_{t-i} + \mu_t \tag{4}$$

Again, one can use an F-test for testing short-run asymmetries ($\psi_1 = \psi_2$).

As extensively argued in Section 2, the Italian dairy market has (potentially) witnessed some structural changes. Accounting for such breaks along the time series data is of great importance though, since structural break may entail different methodological approaches and, primarily, change the economic interpretation of results. Therefore, we hinged on the Zivot-Andrews (1992) test for detecting structural breaks and, whenever a break is found, the Johansen, Mosconi, and Nielsen (2000) cointegration test applied. Indeed, the latter allows for up to two structural breaks into the cointegration relation, with related but different asymptotic distributions to those applying in Johansen canonical tests[27]. If a structural break is detected, the AECM in (3) is then specified as

$$\Delta P_{c,t} = \alpha_3 + \beta_1 \Delta P_{p,t} + (1 - I_t) \cdot (a_1^{+} ECT^{+}_{t-1} + a_2^{-} ECT^{-}_{t-1}) + I_t \cdot (a_3^{+} ECT^{+}_{t-1} + a_4^{-} ECT^{-}_{t-1}) + + \sum_{i=1}^{k-1} \Gamma_j \, \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_j \, \Delta P_{p,t-i} + \mu_t \tag{5}$$

Where $I_t = \begin{cases} 1 & if \quad t \geq T \\ 0 & if \quad t < T \end{cases}$ is the Heaviside indicator function, t is the time-line, and T is the known structural break-point. Such model specification permits one to discern between the two periods before (when $I_t = 0$) and after (when $I_t = 1$) the policy change in terms of price transmission dynamics. Intuitively, the usual F-Test can be deployed for testing asymmetries within and between the two different scenarios.

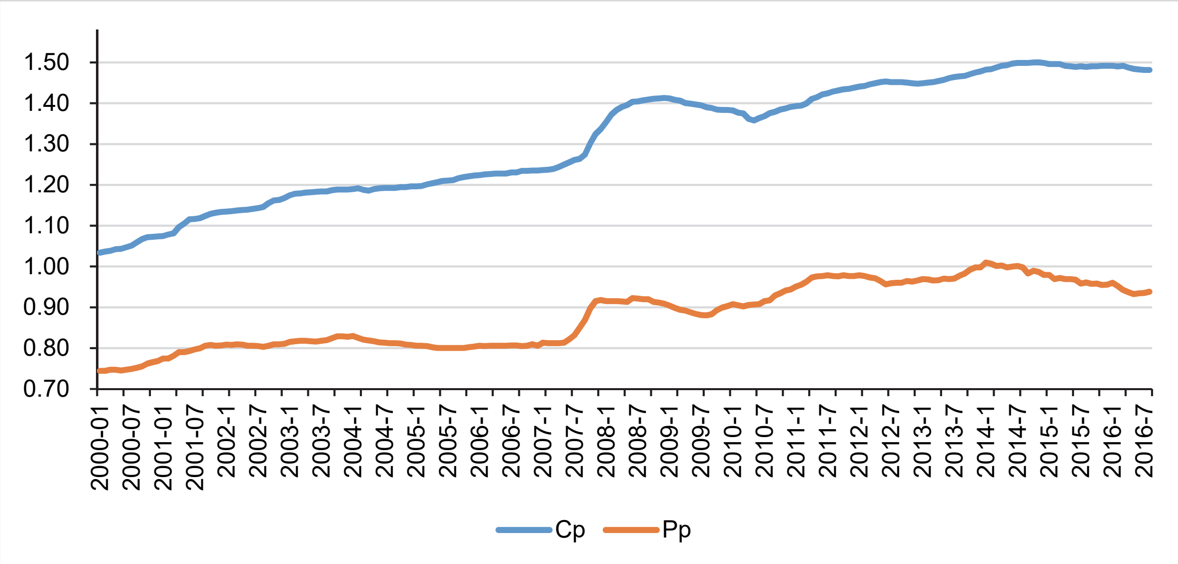
2.5. Data and Results

We used monthly price data at two different levels of the milk supply chain: the price paid by retailers to the industrial milk processor (P_p) and the price applied by the same retailers to the consumers (P_c). One of the largest Italian food distributors provided the authors with monthly prices for fresh milk for the year 2010. Using two price indexes provided by the ISTAT and indexed by the year 2010, time series for the two prices in levels were calculated. The time span covered began in January 2000 to end in August 2016, accounting for 200 observations (see Figure 8). Aimed at mitigating the fluctuations and increasing the likelihood of stationarity after first differencing, we transformed our series into their logarithms. Moreover, this allows for interpreting results in percentage change terms, with

[27] See the Appendix for the model specification

the long-run coefficient β representing the price transmission elasticity (Ben-Kaabia and Gil, 2007; Hamilton, 1994).

Figure 8 – Producer and Consumer Prices of Fresh Fluid Milk on the Italian Market, 2000-2016



Source: Authors own elaboration.

We first test for the presence of a unit-root in our price series. A constant was included since it accounts for the current margin between the two series over time. Several unit-root tests were employed, as the literature suggests, to conclude if the series contains or not a unit-root firmly. The GLS-ADF test (Elliott et al., 1996) and the KPSS test (Kwiatkowski et al., 1992) have been additionally employed to the PP test (Phillips and Perron, 1988) to overcome the low power and size distortions bore by canonical ADF and PP tests (DeJong et al., 1992; Ng and Perron, 2001; Schwert, 1989). As shown in Table 14, all tests lead to the conclusion the two series are I(1) differenced stationary processes[28].

From the graphical inspection and from the events in the economic calendar we described in previous sections, one might suspect that some structural change occurred in 2007. Hence, the Zivot-Andrews test (1992) for detecting structural breaks was applied, spotting a change in August 2007 (see Table 15 for further details).

[28] The same battery of unit-root test was applied to both series differenced one time, leading to the result they are stationary.

Table 14 – Unit-Root Tests for CPI and PPI

Unit-Root Test	Lags	Tau-Stat.	1% C.V.	5% C.V.	10% C.V.	IC	Results
				P_c			
DF-GLS (w/Trend)*	4	-1.859	-3.46	-2.911	-2.625	Ng-Perron	I(1)
	1	-1.125	-3.46	-2.936	-2.647	SBIC, MAIC	
PP*	4	-2.411	-3.477	-2.883	-2.573	Newey-West	I(1)
PP (w/Trend)*	4	-1.178	-4.007	-3.437	-3.137	Newey-West	
KPSS (w/Trend)	4	0.47	0.217	0.148	0.12		I(1)
KPSS	4	3.94	0.739	0.462	0.348		
				P_p			
DF-GLS (w/Trend)*	11	-2.116	-3.46	-2.837	-2.557	Ng-Perron, MAIC	I(1)
	2	-2.162	-3.46	-2.928	-2.64	SBIC	
PP (no Trend)*	4	-1.691	-3.477	-2.883	-2.573	Newey-West	I(1)
PP*	4	-1.248	-4.007	-3.437	-3.137	Newey-West	
KPSS (w/ Trend)	4	0.264	0.217	0.148	0.12		I(1)
KPSS	4	3.727	0.739	0.462	0.348		

***Maximum lag-length selection set to 12**

Source: Authors' calculations

Table 15 - The Zivot-Andrew Test for Structural Break

		P_c		
lags	break	t-stat	10%	Results
1	2007m9	-3.819	-4.58	I(1)
		P_p		
lags	break	t-stat	10%	Results
2	2007m6	-4.167	-4.58	I(1)

Source: Authors' calculations

Accordingly, when testing for cointegration one has to account for the presence of such break, and therefore the standard Johansen test would not be appropriate. The Johansen, Mosconi, and Nielsen (2000) methodology has been used. A constant has been restricted to the cointegrating relationship and results are detailed in Table 16.

Table 16 – Johansen Trace Test for Cointegration

Rank	Trace	Frac95	P-Value
0	27.832	26.406	0.033
1	8.224	12.836	0.266

The number of lags to include was selected accordingly to the SBIC, and it was set to 1

Source: Authors personal calculations

By normalizing on consumer price, we obtain $Cp = 0.206Pp + 0.149D_t + 3.576$, where D_t is the dummy variable for structural break with $\mathrm{t} = 2007{:}08$. Only 20.6% of a one-unit shock in the producer price is transmitted to the consumers. This clearly appoints to an imperfect price transmission, since the relationship between the two prices is not $\beta(1, -1)$. Regarding the adjustment coefficients, we carried out the weak-exogeneity test in order to statistically prove which price is weakly-exogenous (i.e., does not, adjust to the long-run disequilibria). Producer price is weakly-exogenous, confirming the results of previous studies that cost-push mechanism leads the dynamic of PT in agricultural markets (Abdulai, 2002; Ben-Kaabia and Gil, 2007; von Cramon-Taubadel, 1998; Santeramo and von Cramon-Taubadel, 2016)[29]. Moreover, in order to reinforce the exogeneity assumption, Granger-Causality test was investigated, leading to the exogeneity of producer price (see Table 17).

Table 17 - Granger Causality Wald Test for a VAR(2)

Equation	Excluded	χ^2	d.f.	p-value
$\Delta \boldsymbol{Cp}$	ΔPp	18.285	2	0
$\Delta \boldsymbol{Pp}$	ΔCp	2.108	2	0.348

Note the null hypothesis tests if the excluded variable does not Granger cause the independent. A p-value greater than 0.05 (i.e., 5%) means we cannot reject the null and, hence, the variable set as independent is exogenous

Source: Authors own calculations

Before the modeling of APT, one more step is needed, that is the computation of the Error Correction Term; from the linear cointegrating regression, $ECT_t = Cp - \alpha_0 - Pp - D_{2007:08} = \mathrm{Cp} - \alpha_0(-0.004) - \mathrm{Pp}(0.994) - D_{2007:08}(0.040)$. Given the exogeneity of the Pp[30], we set Cp as the independent variable for estimating the (A)ECM.

First, we estimated an ECM to understand if the two prices eventually adjust in the long-run. Results are presented in Table 8 below, and they confirm there exists a long-run equilibrium between producer and consumer prices (the lagged error correction term in the third row is significantly different from 0).

Table 18 – Estimates from the ECM(1)

ΔCPI	Coef.	Std. Err.	t-Stat	P>t
$\Delta \boldsymbol{PPI_{t-1}}$	0.115771	0.034235	3.38	0.001
$\boldsymbol{ECT_{t-1}}$	-0.01554	0.006114	-2.54	0.012
$\Delta \boldsymbol{CPI_{t-1}}$	0.560485	0.057423	9.76	0
$\boldsymbol{\alpha_0}$	0.000651	0.00019	3.44	0.001

Source: Authors own calculation

To deepen the understanding of asymmetric price dynamics in the short-run only, we specified an asymmetric short-run model (ASRM) such as $\Delta P_c = \alpha_0 + \sum_{i=1}^{k-1} \Gamma_i \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_i \Delta^+ Pp_{t-i} + \sum_{i=1}^{k-1} \psi_i \Delta^- Pp_{t-i} + \mu_t$, whose estimates are presented in Table 19.

[29] For the Cp and Pp results of weak-exogeneity test were 10.609 and 3.453, respectively, before a 5% C.V. of 3.841. Therefore, we accept the null of $\alpha_i = 0; i = Cp, Pp$ in the producer price case only.

[30] The estimation of the marginal model suggested in von Cramon-Taubadel (1998) for ensuring the exogeneity is detailed in the Appendix I.

Table 19 – Estimates from the ASRM(1)

ΔCPI	Coef.	Std. Err.	t	P>t
$\Delta^{+}\boldsymbol{PPI_{t-1}}$	0.142415	0.044843	3.18	0.002
$\Delta^{-}\boldsymbol{PPI_{t-1}}$	0.108102	0.07538	1.43	0.153
$\Delta\boldsymbol{CPI_{t-1}}$	0.599437	0.056387	10.63	0
$\boldsymbol{\alpha_0}$	0.000505	0.000239	2.12	0.036

Source: Authors own calculation

Only positive changes in the producer price are significant, and therefore only when producer price increases (i.e., margins are squeezed) the consumer price response is significant. This dynamic seems quite in line with the existing literature claiming prices behave like "feathers" when margins stretch and "rockets" when they squeeze. However, estimating an AECM will provide a deeper understanding of asymmetric price dynamics. We estimate the model expressed in (2), although we excluded the contemporaneous effect of the ΔPp_t, since retail prices in agricultural markets take time to respond to shock, such as $\Delta P_c = \alpha_0 + a_1^{+} ECT_{t-1}^{+} + a_2^{-} ECT_{t-1}^{-} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \, \Delta P_{p,t-i} + \varepsilon_t$.

Table 20 – Estimates from the AECM(1)

ΔCPI	Coef.	Std. Err.	t	P>t
$\Delta\boldsymbol{CPI_{t-1}}$	0.539381	0.057903	9.32	0
$\boldsymbol{ECT_{t-1}^{+}}$	0.003487	0.011152	0.31	0.755
$\boldsymbol{ECT_{t-1}^{-}}$	-0.04197	0.014346	-2.93	0.004
$\Delta\boldsymbol{PPI_{t-1}}$	0.105854	0.034311	3.09	0.002
$\boldsymbol{\alpha_0}$	0.000124	0.000321	0.39	0.7
F-test	$\boldsymbol{H_0}$	**F-stat**	**p-value**	
F(1,194)	$H_0 = ECT_{t-1}^{+} + ECT_{t-1}^{-} = 0$	9.23		0.002

Source: Authors' calculation

Only when $P_p > P_c$ (i.e. negative ECT) the system adjusts to the steady-state (i.e. Positive Asymmetries). This result is consistent with the ASRM, in which only positive changes in the producer price are significant to the consumer price. An F-test was carried out in order to check if there exists asymmetry in the price transmission process and the statistic resulted to be smaller than the C.V. at 1% (see the last row in Table 20), hence we rejected the null of symmetric adjustment. To deepen the understanding of asymmetric price dynamics, an asymmetric short-run model (AECM-SR) as in (3) was estimated, and results are reported in Table 21.

Only positive changes in the producer price are significant, and therefore only when producer price increases (i.e., margins are squeezed) the consumer price response is significant. This dynamic seems quite in line with the existing literature, claiming prices behave like "feathers" when margins stretch and "rockets" when they squeeze. This is confirmed by the behavior of the ECT. Both F-tests (see the last two rows of Table 21) confirm the existence of both long and short-run asymmetries.

Table 21 – Estimates from AECM-SR(1)

ΔCPI	Coef.	Std. Err.	t	P>t
$\boldsymbol{\Delta CPI_{t-1}}$	0.534472	0.058588	9.12	0
$\boldsymbol{ECT^{+}_{t-1}}$	0.003394	0.011172	0.3	0.762
$\boldsymbol{ECT^{-}_{t-1}}$	-0.04242	0.01439	-2.95	0.004
$\boldsymbol{\Delta PPI^{+}_{t-1}}$	0.12241	0.044271	2.77	0.006
$\boldsymbol{\Delta PPI^{-}_{t-1}}$	0.066435	0.074801	0.89	0.376
$\boldsymbol{\alpha_0}$	0.000036	0.000353	0.1	0.919
F-test	$\boldsymbol{H_0}$	**F-stat**	**p-value**	
F(1,194)	$ECT^{+}_{t-1} + ECT^{-}_{t-1} = 0$	9.41	0.0025	
F(1,194)	$\Delta^{+}Pp_{t-1} - \Delta^{-}Pp_{t-1} = 0$	5.74	0.0175	

Source: Authors' calculation

However, we might introduce the structural break we found in the initial steps of our analysis. Therefore, we estimated the following AECM-SR model, which accounts for a break in the ECT:

$$\Delta P_{c,t} = \alpha_3 + \beta_1 \Delta P_{p,t} + (1 - I_t) \cdot (a_1^{+} ECT^{+}_{t-1} + a_2^{-} ECT^{-}_{t-1}) + I_t \cdot (a_3^{+} ECT^{+}_{t-1} + a_4^{-} ECT^{-}_{t-1}) +$$
$$+ \sum_{i=1}^{k-1} \Gamma_j \, \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_j \, \Delta P_{p,t-i} + \sum_{i=1}^{k-1} \psi_1 \, \Delta^{+} Pp_{t-i} + \sum_{i=1}^{k-1} \psi_2 \, \Delta^{-} Pp_{t-i} + \mu_t$$

Estimates (see Table 22) confirmed the previous results, even though after the structural break the term $ECT^{+}_{D,t-1}$ turns significatively different from zero, meaning that when margins are stretched, consumer price re-adjusts to the equilibrium also. In order to understand if there still exists some asymmetries in price dynamics, the F-Test (see the last row of Table 22) accepts the null of equality. Therefore, in the post-break period the price transmission turned (more) efficient and become symmetric.

Table 22 – Estimates from the AECM-SR(1) with a structural break in the ECT

ΔCPI	Coef.	Std. Err.	t	P>t
$\boldsymbol{\Delta CPI_{t-1}}$	0.494913	0.060311	8.21	0
$\boldsymbol{ECT^{+}_{t-1}}$	0.015428	0.012174	1.27	0.207
$\boldsymbol{ECT^{-}_{t-1}}$	-0.03901	0.01485	-2.63	0.009
$\boldsymbol{ECT^{+}_{D,t-1}}$	-0.03723	0.017131	-2.17	0.031
$\boldsymbol{ECT^{-}_{D,t-1}}$	-0.01784	0.017358	-1.03	0.305
$\boldsymbol{\Delta PPI^{+}_{t-1}}$	0.125666	0.045739	2.75	0.007
$\boldsymbol{\Delta PPI^{-}_{t-1}}$	0.005552	0.080136	0.07	0.945
$\boldsymbol{\alpha_0}$	1.68E-05	0.000351	0.05	0.962
F-test	$\boldsymbol{H_0}$	**F-stat**	**p-value**	
F(1,190)	$ECT^{+}_{D,t-1} + ECT^{-}_{D,t-1} = 0$	5.17	0.024	
F(1,190)	$ECT^{+}_{D,t-1} - ECT^{-}_{t-1} = 0$	0.01	0.94	

Source: Authors' calculation

Following Ben-Kaabia et al. (2005) and Santeramo (2015), we estimated the half-lives for each model specification, that is the periods required for the system to achieve $\varepsilon\%$ adjustment to their new equilibrium after an exogenous shock occurred. They are expressed at the same time series frequency of the data one has used. For an ECM specification, they are expressed as $H = \frac{\ln(1-\varepsilon)}{\ln(1+\rho)}$, where ε is the factor of adjustment and ρ is the speed of adjustment of the ECT, i.e. the, associated coefficient (α_j), and H is the number of months the system implies to reach the equilibrium again.

Table 23 – Half-lives for each model specification

Model	Coefficient	% ε	ρ	H
AECM	ECT^{-}_{t-1}	0.5	-0.04197	16.166
		0.75		32.333
		0.9		53.703
		0.99		107.41
AECM-SR	ECT^{-}_{t-1}	0.5	-0.04242	15.991
		0.75		31.982
		0.9		53.121
		0.99		106.24
AECM-SR (w/ structural break)	ECT^{-}_{t-1}	0.5	-0.03901	17.42
		0.75		34.839
		0.9		57.867
		0.99		115.73
	$ECT^{+}_{D,t-1}$	0.5	-0.03723	18.269
		0.75		36.538
		0.9		60.689
		0.99		121.38

Source: Authors' calculations

Looking at the 90% adjustment, there is a slight decrease in the speed of adjustment in the post-break scenario: when margin is squeezed the system spends 9% more time to go back to the steady state (i.e., roughly 4 months more); moreover, in a stretching-margin scenario, it employs 14% more time to adjust (i.e., about 7 more months).

2.6. Conclusions

The paper explores the VPT process within the Italian dairy industry, a major agrifood sector that has been on the spot in recent years because of the liberalization policy put in place by the EU. This work contributes to that branch of literature analyzing the CAP's effects on food supply chains. Firstly, we found a cost-push transmission, that is going from producers to retailers only (see Santeramo and von Cramon-Taubadel (2016) for an exhaustive list of work supporting this causal relation). We then assessed the existence of positive APT in the Italian dairy industry, and, accordingly, only when margins are squeezed the consumer price responds to shocks. This behavior is consistent with Kinnucan and Forker (1987) study on dairy products in the U.S., stating that in the case of government interventions (i.e., price support for farmers) the PT process may result to be asymmetric. Once determined the structural break in August 2007, results showed that in the post-break period the PT dynamics turned symmetric. Relaxing the government (i.e., the EU in this case) intervention on the milk market seems brought it closer to its natural functioning. Indeed, giving the high perishability of the fresh milk, symmetric PT is consistent with recent literature findings: Serra and Goodwin (2003) found symmetric dynamics for perishable dairy products (i.e., fresh milk), whereas Santeramo and von Cramon-Taubadel (2016) speaks in favour of symmetric PT for high-perishable products - in their analysis, the authors found out that asymmetries for those products with a lower shelf-life is quite a rare phenomenon, since over 16 investigated commodities only 1 resulted in asymmetric behaviour. Moreover, Heien (1980) stated how, due to their dynamic price behavior, changing price for perishable products is less of a problem. On the other hand, part of the recent literature suggests that when high-perishable products are considered, negative asymmetries arise (see Ward (1982), Kim and Ward (2013) and Santeramo (2015)). Cacchiarelli et al. (2016) investigated the impact of CAP liberalization reforms on the pasta-chain in Italy, and despite their results are not directly comparable with our findings regarding product's, chain, and policy characteristics, they draw similar conclusions. Considering the farm-wholesaler relationship, they found the liberalization process turned the price adjustment symmetric; on the opposite, when the relationship wholesaler-retailer is concerned, the PT became asymmetric in the liberalized scenario. Although the policy change has turned the milk market more efficient, half-lives suggest that the system's response to shocks became more sluggish. The same result is found in Lee and Gómez (2013) in their study on the impact of the end of coffee quota exports on PT dynamics in international markets: the increase in volatility (a direct consequence of a more free market), generate more uncertainty since the domestic milk price is now more integrated with the world market. Retailers are then more reluctant to change prices, facing high menu costs. The increase in volumes of trade may have also modified the slopes of the consumer demand curve as well as cross-elasticities with new milk substitutes. In the last decade, consumer preferences have changed, cow fluid milk consumption decreased, and consumer search costs plunged. Consequently, the more the consumers became sensitive to price changes, the more the retailers have to accommodate their strategy to the market course, fearing spoilage (Santeramo and von Cramon-Taubadel, 2016). Finally, when margins are increasing the system adjusts to the equilibrium at a slower pace on the opposite situation, suggesting retailers maintaining some market power enlarging periods of positive surplus. Of course, results are far from being conclusive and have to be taken with caution. Especially for perishable products, weekly data are more appropriate since lower frequencies (e.g., monthly) may not capture rapid price movements (Santeramo, 2015). Moreover, the present study does not consider the farm-gate level, which is of high interest when dealing with CAP reforms. The latter surely suggest a fundamental future path of research.

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3. Price Transmission Analysis in the Italian Feed Industry: an Interrupted Threshold-Cointegration Approach

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Abstract

We investigate the vertical price transmission (VPT) mechanism in the Italian milk cow-related feed industry, especially the interdependence between Italian corn and feed prices. We propose a novel two-regimes threshold-cointegration model, where regimes are triggered by an observable transition variable. The latter features a wide range of both fundamentals and non-fundamentals variables that recent literature put on the spot as potential drivers for the surge in agricultural commodity prices occurred in 2007. Abandoning the cemented assumption of a continuous cointegrating relationship, we allow for interruptions in the error correction mechanism, capturing which economic drivers impact the VPT process. Empirical results suggest the impact of non-fundamentals is negligible, whereas market fundamentals still play a significant role in shaping price transmission dynamics.

3.1. Introduction

The breakout of the world economic crisis in 2007 and the related commodity price boom opened the floor to a broad range of studies investigating the causes of such rise in commodity prices. Agricultural commodities have been on the spot, especially cereals and protein crops. Scholars often charged non-fundamentals drivers (i.e., energy-related prices and financialization of agricultural commodities) as the source of such price peaks (see among others Baffes and Haniotis, 2016; Brümmer et al., 2016; Cooke and Robles, 2009; Han et al., 2015; McPhail et al., 2012; Mensi et al., 2014; Tadasse et al., 2016). However, up to the authors' knowledge, no attempts have been made for disentangling the effect of those drivers on vertical price transmission processes, and the Italian feed industry represents a compelling case-study to investigate. Its revenue represents 8% of the overall agrifood industry, while feed costs account for more than 50% of the overall operational expenses for an average Italian milk producer. Maize is the primary cereal in feed formula, representing 60% of the overall raw inputs. The industry self-sufficiency rate is decreasing, approaching 62% nowadays, enhancing the importance of import prices (i.e., maize and soybean mainly) in shaping VPT dynamics. International trade in agricultural commodities could expose the domestic feed supply chain to exogenous shocks, such as biofuel and financial derivatives, which have been proved to exert significant effects on agricultural prices. On the one hand, according to the last OECD/FAO Agricultural Outlook (2016), the EU is not a significant producer of ethanol[31], whereas it accrues for 40% of the global biodiesel output (authors' calculations based on OECD and FAO, 2016), representing the largest world producer. The latter principally hinges on rapeseed (FEFAC, 2017), and according to last data available from the EBB (European Biodiesel Board, 2013), Italy is the sixth European producer (3.7% of total EU production)[32]. Latest estimates agree the biodiesel production will increase by 45% in the next fifteen years (Avalos, 2014). The openness of the EU cereal markets may entail contagion from larger cereal markets (i.e., U.S.) and its derivatives, although Italian imports refer for more than 95% to European and Ukranian maize. Market fundamentals (i.e., Italian feed and corn prices, together with substitute agricultural commodities and the stock-to-disappearance ratio) are also considered for deepen the understanding of their impacts. To assess the effects of the mentioned drivers on the PT dynamic, we propose a new time series econometrics approach allowing for interruptions in the co-movement between prices, depending on the specific trigger variable under consideration. Indeed, standard price transmission analysis relies on the assumption of a continuous cointegrating relationship. The contribution of this research is three-fold. Firstly, we investigate a crucial agrifood sector that did not receive so far significant attention in the PT literature. Secondly, we disentangle if and how a broad range of market (non)fundamentals variables exert some impact on the VPT dynamics, contributing to the current debate over the so-called "Masters Hypothesis" (Irwin and Sanders, 2012).

3.2. The Italian Feed Industry

Italian feed industry represents a vital sector within the agrofood industry, with an estimated turnover in 2016 of about 6 billion euros and a whole output of more than 14 million tonnes of compound feed, relying upon 429 production units (Assalzoo, 2017). The Regulation (EC), No. 767/2009 of 13 July 2009, defines "compound feed" as "[…] a mixture of at least two feed materials […] in the form of complete or complementary feed" (see Art. 3, letter h), and is mostly the primary feedstuff used for dairy cattle alimentation, the latter representing 23% of full feed production (Assalzoo, 2017). Within EU-28, Italy represents the 6th feed producer, accruing for more than 10% of total feed and 8% of entire cattle feed (data from FEFAC - European Feed Manufacturers' Federation, 2016)[33]. The EU livestock is the most important outlet for EU-cereals since 60% of total production goes for cattle feeding

[31] Ethanol production from maize represents, in 2016, 3.8% of world production, with an estimated increase in 2025 up to 7% of global production.
[32] Germany, France, and the Netherlands accrue for 55% of total production with 24.3%, 18.2%, and 12% respectively.
[33] The first five feed producers are Germany (23.7 Mio tonnes), Spain (22.1 Mio tonnes), France (20.3 Mio tonnes), UK (15.6 Mio tonnes), and the Netherlands (15.6 Mio tonnes).

(FEFAC, 2017). Cereals and oilseed meals are the two primary inputs representing 50% and 27% of the formula composition, respectively (FEFAC, 2017). Corn represents 60% of the cereal category (ISMEA, 2013), and 74% of Italian maize is destined to compound feed production (Assalzoo, 2017; ISTAT, 2017). Considering also the corn directly used as simple feedstuff, then 90% of harvested maize in Italy is destined to cattle feeding (USDA, 2012). The northern Italian regions of Emilia-Romagna, Lombardy, Veneto, and Piedmont accrue for 63% of both volume and feed production units - Emilia-Romagna hosting alone 30% of units (Assalzoo, 2017). Such geographical concentration is explained by the fact that those regions gather 80% of corn-farmers and 82% of national maize production (data from ISTAT, 2017). In its recent study, the European Commission (2016) estimates that for the average Italian milk-producing unit 52% of total operational costs are represented by purchased compound feedstuffs (i.e., 114 €/tonne over 221 €/tonne of produced milk), being the share of the home-grown feed on total feeding cost just 15%. In the earliest 2000s, Italian corn supply almost completely satisfied the feed industry demand, while nowadays the self-sufficiency rate plunged to 62% (see Table 24)[34]. As a consequence, maize imports skyrocketed (+685% for the period 2000-2016), reaching 4.4 Mio tonnes in 2016 (data retrieved from ISTAT, 2017).

Table 24 – Italian Maize Balance Sheet, 2000-2016

Maize (.000 t)	**2000**	**2005**	**2007**	**2008**	**2010**	**2015**	**2016**
Harvested Production (P)	10,140	10,510	9,809	9,723	8,608	7,074	6,840
Imports (M)	477	1,224	2,484	2,200	2,078	3,752	4,351
Exports (X)	181	29	149	116	137	130	120
Apparent Consumption (P+M-X)	10,435	11,705	12,145	11,806	10,550	10,696	11,071
For Feed Production (F)	*7,800*	*9,090*	*9,100*	*8,700*	*8,450*	*7,889*	*8,192*
For Other Uses	*2,635*	*2,615*	*3,045*	*3,106*	*2,100*	*2,807*	*2,878*
Self-Sufficiency Rate [P/(P+M-X)]	97%	90%	81%	82%	82%	66%	62%
Feed - Total Corn Rate [F/(P+M+X)]	75%	78%	75%	74%	80%	74%	74%

Source: Authors' elaboration on data from ISTAT and Assalzoo

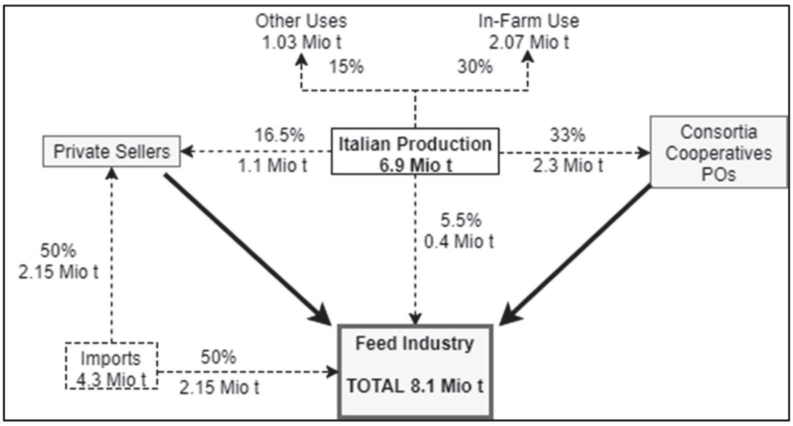
Most of the imports originated from within-EU, with percentages going from 99% in 2000, to 62% in 2016, given the entry of Ukraine (29% of total imports in 2016)[35]. France always played a preponderant role, despite its decreasing importance to the benefit of Ukraine, representing nowadays 9% of total imports – up to 2008 it was one of two most significant import markets, with a market share of 90% in the earliest 2000s.

[34] Both corn' surface and harvest featured a tremendous decreased in the last 17 years (-38% and -33% for the period 2000-2017, respectively) (data retrieved from ISTAT, 2017).

[35] Ukraine cemented its role as the largest corn importer during the last six campaigns, representing 30% of total corn imports. Hungary has maintained its significant role, being since 2005 between the first two most prominernt players, with a market share wandering around an average of 25% with peaks of 45% in certain campaigns. Finally, Austria also plays an important role, with a stable 10% of total imports over the period 2000-2016, with peaks of 15%.

The feed supply chain is composed of three main steps (see Figure 9). The first step (not considering the seed supplier) is the agricultural production, with 133 thousand corn producers distributed on a surface of 650 thousand hectares. According to ISMEA, one-third of maize volume is used in farm, mainly for feeding animals, while around 15% is destined to the milling industry. The harvested maize is then sold to operators who store the production and eventually re-sell[36] it to feed producers. Private buyers trade one-third of the corn availability and half of total imported quantities[37], whereas consortia, cooperatives, and producers' organizations (POs) represent the lion share regarding collected volume, buying 60% of disposable maize. The feed industry buys 10% of corn volume directly from farmers and half of total imports.

Figure 9 – The Italian Feed Supply Chain, 2016



Source: Authors' elaboration on ISMEA, Assalzoo, and ISTAT

According to the last data available, the structure of the Italian feed industry appears quite concentrated, with the first six feed producers gather 42% of total volume and value. The first company (Veronesi) represents alone a market share of 21% (Mambriani, 2009), being the 8th largest feed producer in Europe with a final output of more than 3 million tonnes (Wessler et al., 2015). Following a more general European trend, Italian feed industry is witnessing a decrease in units of production[38] although a stable volume of production, pointing to an increasing concentration (Wessler et al., 2015).

[36] For a thorough description of contracts typology existing between the three steps see Serra and Zuppiroli (2009).
[37] Here we assumed the whole imports are destined to the feed industry.
[38] The number of units decreased by -40% for the period 2003-2016 (data retrieved from Assalzoo).

3.3. Price Transmission in the Feed Market: a Review of the Existing Literature

In Section 3.2 we highlighted the importance of cereals, particularly maize, as the main component of compound feed, especially that for dairy cows. The latter represents the most weighty cost for milk producers. Both price transmission (PT) and price volatility (PV) literature offer a large strand of works accounting for cereals, oil crops, and food prices, but very few directly investigate the spot feed price. Up to the authors' knowledge, this is the first attempt trying to disentangling how main economic triggers blamed for 2006-2008 commodity price jump may affect vertical price linkages along agrifood supply chains.

Agricultural price movements have many implications, from food security to agricultural income. Understanding which are the main triggers affecting the transmission dynamics may help in design strategies to cope with extraordinary market events. The 2006-08 surge in agricultural commodity prices opened the floor to a very large number of works devoted to understand the cause of such escalation, since grain prices increased by a factor larger than two, being among the most substantial jump in commodity price history (Cooke and Robles, 2009; Mensi et al., 2014). The focus has been put mainly on the impacts of energy-related markets, given the upsurge of biofuels production that may have strengthened such linkage. Furthermore, the so-called 'financialisation' of agricultural commodities may enhance spillover effects between derivatives (i.e., futures) and spot markets, and tightening the bond between food and energy too (see among others Baffes and Haniotis, 2016; Brümmer et al., 2016; Cooke and Robles, 2009; Han et al., 2015; McPhail et al., 2012; Mensi et al., 2014; Tadasse et al., 2016). However, the debate among scholars is yet to be solved, due to controversial results. As raised by Beckman et al. (2011, 2012) and Myers et al. (2014), the role that livestock feeds play in the biofuel debate is crucial, since they compete for feed grains, and they are primary users of biofuel coproducts (i.e., the so-called DDGS - dried distillers grains with solubles from corn, and soybeans meal) (Ubilava, 2012).

In a recent paper, Baffes and Haniotis (2016) consider what drivers may explain the agricultural price cycle, arguing against uni-dimensional approaches that put too much weight on a single factor. Investigating five agricultural commodities (i.e., maize, soybeans, wheat, rice, and palm), they conclude that energy prices (i.e., U.S. retail oil price) entailing the most significant effect on agricultural prices, followed by the stock-to-use ratio. The latter can be interpreted as a proxy for biofuels production, as decreasing stocks may be due to feedstock usage in producing energy commodities - the shrinkage that maize stocks experienced in the period, 2005-2014 has been explained by the increasing biofuel production[39]. Avalos (2014) also investigates the relationship between corn, soybeans, and oil prices, pointing out how the from 2006 a new long-run relationship between maize and crude oil began to exist, and how in the short-run energy prices affects both maize and soybeans. The use of ethanol for energy purposes is strictly dependent on the price of oil relative to ethanol, strengthening the fuel-food linkage[40]. Baffes (2013) states energy, crude oil in particular, affects food price formation, and how the latter will play a pivotal role in the next future, not only for its importance in production costs but also because of its substitutability with biofuels. Wang et al. (2014) discern between the nature of oil price shocks effects on agricultural commodities price, concluding that oil-specific demand shock in the post-crisis period (2006-2008) turned highly significant. Being those defined by future expectations of oil supply shortfall and speculative demand, they highlight the importance of considering speculation as a source of food price spikes. Chang and Su (2010) found significant spillover effects from oil to corn and soybean futures during higher oil price windows, concluding the substitution effect between fossil and biofuels causes food price spikes. Likewise, Natanelov et al. (2011) results confirm that cointegration between U.S. oil and corn futures only exists when crude oil trespasses the 75$/barrel threshold, with biofuel US policy measures buffering the linkage between food and fuel. Mensi et al. (2014), using both American and European oil prices, together with American FOB cereal prices, found the linkage oil-food significant and affected by OPEC announcements. Koirala et al. (2015) investigate the relationship between (high-frequency) different energy and agricultural futures, finding a positive and significant correlation between the twos. Nicola et al. (2016) and specify that the degree of co-movement in recent years between energy and agricultural price returns, especially between maize and soybean oil, augmented. They argue this could be a biofuels effect since those are the two primary inputs for biofuels production in the U.S. Both Kristoufek et al. (2012) and Tyner (2010) found that when food prices are low, the connection of agricultural commodities with energy prices is weak, whereas high food price regimes (i.e., post-crisis period) trigger a stronger linkage, especially between ethanol and corn. Concerning biodiesel, this does not show such high bound, highlighting the importance of the considered type of biofuel. Not much attention has been devoted to the European markets. Peri and Baldi (2010, 2013) consider the EU biodiesel scenario, relying on two price series, diesel and biodiesel prices[41]. They conclude

[39] Note that since 2005 the Energy Policy Act (EPA) allows ethanol to be the only additive for gasoline to comply with new standards (Avalos, 2014). U.S. corn is the main source for ethanol production, and the U.S. is the largest biofuel producer.
[40] That is to say, a shock to the oil price is exogenous and transmitted to food commodity prices via the (endogenous) production and consumption of biofuels (Avalos, 2014).
[41] For biodiesel price they used as a proxy the Rapeseed oil FOB price at ARA. They also found strong correlation between the latter and the two series Rapeseed Methyl Ester and the Fatty Acid Methyl Ester.

saying that EU policy[42] eventually generated a new relationship between fossil fuel and vegetable oil prices, with rapeseed oil price shifting from its own-market to the diesel market. Abdelradi and Serra (2015) found bidirectional volatility asymmetric spillovers between Spanish biodiesel and sunflower oil prices, while Busse et al. (2012) conclude that a stable long-run relationship exists between German biodiesel and rapeseed and soybean oil prices.

On the other hand, many empirical works point to the opposite. Cooke and Robles (2009), investigating international prices of corn, wheat, rice, and soybeans, found the only significant element in causing rising food prices is financialization and speculation, whereas oil and biodiesel resulted in weak importance in explaining increasing food prices. Myers et al. (2014) conclude co-movement between energy and U.S. agricultural feedstock prices (i.e., corn and soybeans) do not share a tight connection. Although such connection exists, it tends to vanish in the long-run, meaning they do not share a common stochastic trend, so they wander away one from another. Saghaian (2010) neither conclude a causal link exists between oil and agricultural commodity markets. Zhang et al. (2010) studying U.S. cereal and fuels prices, found no long-run relation together with fragile short-run relations. Reboredo (2012) found no-contagion between international crude oil and food prices, supporting a sort of neutrality of agricultural markets to changes in oil prices. Zhang et al. (2009) neither can conclude a long-run relation exists, in recent years, among fuels (ethanol and crude oil) and agricultural commodities (corn, and soybeans) prices. On the European side, especially regarding the Italian market, Serra and Zilberman (2013) provide a thorough literature review of biofuel-related price transmission works, indicating how for Italy the oil price does not drive feedstock prices. Furthermore, Serra (2013) states how changes in biofuel prices may not have a substantial impact on food prices not because they are not determinants, but rather because they affect it through volatility spillovers. López Cabrera and Schulz (2016) investigated European future prices for rapeseed oil, crude oil, and German biodiesel, and a stable long-run equilibrium relationship between the agricultural commodity and biodiesel exists, with both prices reacting to changes. However, biodiesel does not influence rapeseed oil nor crude oil prices in the short-run, with no volatility linkages between the three prices. Therefore, they conclude concerns about biodiesel increasing food price is not founded.

[42] The EU policy is intended to limit the impact of increases in oil price, achieving fuel security and reduce GHG emissions. Specifically, see the EU Council Directive 2003/30/EC and 2009/28/EC. As indicated in (Peri and Baldi, 2013), both the blending mandate and the tax relief have been the main triggers for biodiesel production expansion in the last years.

According to the ‘Masterson Hypothesis,' defined by Irwin and Sanders (2012), index investors activity generated massive bubbles in both food and energy markets[43]. Is this the so-called ‘financialization’ of commodity markets, that is agricultural commodities started to be considered as an asset by financial investors (e.g., banks, speculators, equity found among others). Whether this has its positive points – in the long-run, it should improve both risk sharing and hedging potential, together with price discovery and stability – short-term dynamics may hamper the market functioning, weakening the connection between price and fundamentals (Venables, 2014). Tang and Xiong (2012) spot the early of the 2000s as the key-moment for such financial dynamics[44] in agricultural markets, given the dramatic growth of commodity index investments. They found a correlation between energy and non-energy commodities futures increased from 2004, arguing that the negative side of such financialization being the volatility spillover effects this may generate from outside-markets (i.e., energy) to and across commodities. Gilbert (2010) results also suggest that financialization is behind the linkage oil-food via index futures investment, although correlation “is the result of common causation and not of a direct causal link” (p. 420). McPhail et al. (2012) investigate corn price volatility in the U.S., concluding that in the short run speculation plays a major role, whereas energy prices drive volatility in the long-run. Gilbert and Pfuderer (2014) concluded there is a causal relationship binding CIT (Commodity Index Traders) and agricultural commodity future price returns, but that is not structural and informational only. They also conclude that generally speaking, there is weak evidence about how trading indexes directly affect agricultural prices. Indeed, Irwin (2013) after an extensive literature review concluded that Masters Hypothesis is invalid, since not soundly supported by empirical evidence. The recent work of Etienne et al. (2015) also supports this charge, besides stating that agricultural price explosive bubbles are exceptional, short-lived, and small in magnitude. Likewise, the role of futures is of importance here. Nowadays, traders can access price information for different commodities in different parts of the world easily and quickly, without any formal trade between the two investigated places, and futures are derivative markets allowing for price dissemination[45]. As pointed by Hernandez and Torero (2010), spot prices are discovered in futures, so that changes in future prices affect agricultural prices on physical markets. Ganneval (2016) specified a model in which is the flow of information rather than physical trade leading to spatial price transmission, using the level of volatility as the trigger for different agricultural spot markets linkages. He concludes that for the European market of corn, feed barley, and

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[43] For further details, see the two Masterson’s testimonies before the Committee on Homeland Security and Governmental Affairs in 2008, and the Commodity Futures Trading Commission in 2009.
[44] More specifically, they specify that consequent to the collapse of equity market in 2000, “[…] the widely publicized discovery of a negative correlation between commodity returns and stock returns” triggered such volume of investments on commodities (Tang and Xiong, 2012, p. 56).
[45] See Hayenga and Schrader (1980) about formula pricing.

protein pea, in periods of high volatility, the system converges quickly to its equilibrium. In other words, deviations from long-run steady –state are corrected faster, and price adjustment after a shock takes less time. Nicola et al. (2016) found stock market volatility positively correlated with co-movements of price returns across markets. Future prices incorporate all the available information, so they can have a role in tightening or weakening vertical price linkages.

3.4. Methodology and Data

Time-series econometric approaches are non-structural models aimed at identifying empirical (ir)regularities in the price series. Because of their low-demanding volume of data, they represent an agile and cemented methodology for characterizing price behavior, shedding light on economic indicators' patterns (Serra, 2013). The lion share of works on PT hinges on (non)linear cointegration methods (see among others Ben-Kaabia and Gil, 2007; Ben-Kaabia et al., 2005; Hassouneh et al., 2010). Given the use of a varied set of variables, threshold modeling technique is approached, given that it allows testing for non-linear adjustments in the PT dynamics. Whenever the standard Engle and Granger (1987) cointegration procedure does not reject the null of no-cointegration between feed and corn prices, further non-linear approaches such those of Balke and Fomby (1997) and Hansen and Seo (2002) cannot be applied. However, prices may co-move only in specific periods, due to sudden market changes. That is to say, they share a common stochastic trend according to specific periods, whereas classical Error Correction Model (ECM) technique assumes by definition a continuous cointegrating relationship. Trying to overcome such drawback, we propose a novel modeling technique, generalizing the approach of Balke and Fomby (1997). We developed a two-regime cointegration model where the regimes are selected by an observable transition variable (the "trigger"), starting from the time series econometric model offered in Psaradakis et al. (2004) and Martins and Gabriel (2014). This allows for capturing which market-drivers may impact the price transmission process since the transition variable is used for discovering the origin of changes in the PT dynamics between the feed and corn.

In their seminal paper, Balke and Fomby (1997) developed a two-step approach in which they tested for linear cointegration – following Engle-Granger- for later testing for the presence of non-linearity against the null of linear cointegration. We propose two developments in the context of threshold cointegration: first, we test for threshold cointegration without testing first for standard linear cointegration; second, we consider the case in which the equilibrium term is characterized by non-stationary behaviour in one regime, as in Psaradakis et al. (2004), and Martins and Gabriel (2014). We consider two prices, x_t (i.e., the corn price) and y_t (i.e., the feed price), characterized by the following cointegration relation:

$$\begin{cases} y_t = \mu + \beta x_t + z_t, & z_t = G(s_{t-d})z_{t-1} + \epsilon_t \\ y_t = \gamma x_t + B_t, & B_t = B_{t-1} + \eta_t \end{cases} \tag{5}$$

where s_{t-d} is a stationary transition variable and assuming the equilibrium term z_t follows a TAR model of order one, that is

$$G(s_{t-d}) = \left(\delta_1 I_{\{s_{t-d} \leq c\}} + \delta_2 I_{\{s_{t-d} > c\}}\right). \tag{6}$$

Cointegration between x_t and y_t exists only if the equilibrium error z_t follows a stationary process. The condition of strict stationarity[46] can be satisfied even if one of the two regimes is non-stationary, that is if ($\delta_1 = 1,\ |\delta_2| < 1$) or ($\delta_2 = 1,\ |\delta_1| < 1$)[47]. Therefore, z_t follows a 'globally' stationary process, despite the presence of local non-stationary behaviors, and according to Psaradakis et al. (2004), and Martins and Gabriel (2014), movements towards the long-run equilibrium are not always present. In fact, depending on the value of s_{t-d}, z_t can be locally stationary or non-stationary. Parameters δ_1 and δ_2 determine the speed of the mean reversion for the first and second regime, respectively, as the half-life[48] can be obtained as $h_i = \frac{\log(0.5)}{\log(\delta_i)}$.

A simple two-step procedure is adopted for estimating (1) and (2): in the first step, β, and μ are estimated via OLS, obtaining the residuals $\hat{z}_t$[49]; the second step follows Caner and Hansen (2001), hence we use concentrated OLS for $\hat{z_t} = \left(\delta_1 I_{\{s_{t-d} \leq c\}} + \delta_2 I_{\{s_{t-d} > c\}}\right)\hat{z}_{t-1} + \epsilon_t$ in order to estimate c, δ_1, δ_2[50]. In a vector error correction form, (1) and (2) can be specified as

$$\begin{cases} \Delta x_t = \alpha_x z_{t-1}(G(s_{t-d}) - 1) + \nu_{x,t} \\ \Delta y_t = \alpha_y z_{t-1}(G(s_{t-d}) - 1) + \nu_{y,t} \end{cases} \tag{7}$$

where $\alpha_x = 1/(\gamma - \beta)$, $\nu_{x,t} = (\epsilon_t - \eta_t)/(\gamma - \beta)$, $\alpha_y = \gamma/(\gamma - \beta)$ and $\nu_{y,t} = (\gamma\epsilon_t - \beta\eta_t)/(\gamma - \beta)$. From (3) is clear how the error-correction mechanism is present only when $G(s_{t-d}) \neq 1$.

Non-linearity and cointegration tests are needed for testing assumptions in model (1) and (2), and both were designed on Caner and Hansen (2001) work, with critical values obtained

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[46] Strict-Stationarity condition: $E\left[\left(G(s_{t-d})\right)^2\right] < 1$ (González and Gonzalo, 1997). Further conditions to be satisfied in order ensure stationary are $P[s_{t-d} > c] > 0$, $P[s_{t-d} \leq c] > 0$, $E[\max(0, \log(\epsilon_1))] < \infty$, $essential\ supremum(\epsilon_1) < \infty$.

[47] Under this assumption it is easy to show that $E\left[\left(G(s_{t-d})\right)^2\right] = \delta_1^2\ P[s_{t-d} \leq c] + \delta_2^2\ P[s_{t-d} > c] < 1$.

[48] The half-life details the interval that the process (in regime i) needs to correct 50% of the shock that puts the system out of the equilibrium.

[49] Note that when z_t satisfies the 'α mixing' conditions (Assumption 2.1 in Phillips (1987)), the least squares estimate of β is super-consist (N. S. Balke and Fomby, 1997). Moreover, a sufficient conditions for TAR models with one non-stationary regime to be 'α mixing', is the Markovianity of the transition variable s_t (Gonzalez and Gonzalo, 1997).

[50] The parameter d is usually selected 'a priori', e.g. $d = 1$.

via the bootstrap procedure[51]. We propose a simple two-stage testing procedure: In the first step, following Engle and Granger (1987), we obtaining the error correction terms $\hat{z}_t$ by standard OLS; in the second step, we do not implement the standard unit root tests (such as the Augmented Dickey-Fuller (ADF) test), since $\hat{z}_t$ does not follow a simple autoregressive process. Therefore, we adopt unit root tests for which the alternative hypothesis is a two-regime TAR with (potentially) an autoregressive unit root. Caner and Hansen (2001) developed the asymptotic theory for inference on unrestricted two-regime TAR with autoregressive unit root. They showed that their tests are more powerful than the standard ADF test when the data-generating-process is a TAR processes with local unit roots. Compared to tests developed by Balke and Fomby (1997) and Hansen and Seo (2002), they do not suffer from the reduced power nested in unit-root tests when non-linearity is present in the data-generating process. Moreover, such drawback is amplified in a two-regimes TAR with one non-stationary regime (Caner and Hansen, 2001; Gonzalez and Gonzalo, 1997). Concerning non-linearity test, (1) and (2) becomes a linear model under the null hypothesis (H_0: $\delta_1 = \delta_2 = \tilde{\delta}$), and the standard Wald statistics is employed[52]. Bootstrap approximation has been used for the asymptotic distribution of W, specifying p-values and critical values through the following steps[53]:

i) Parameters estimation under the null H_0: $\delta_1 = \delta_2 = \tilde{\delta}$;

ii) Generation of $N = 20{,}000$ replications - with bootstrap re-sampling techniques – for $\{x_t^*, y_t^*\}_n$, $t = 1, \dots, T$, $n = 1, \dots, N$, using the parameters estimated in i);

iii) Calculation of the test statistic W_n for each series $\{x_t^*, y_t^*\}_n$;

iv) Calculation of the q% critical value of W from the q-quantile of the distributions of the test statistics $\{W_n, n = 1, \dots, N\}$. The p-value of W is given by the frequency of test statistic W_n larger than a given value of W, that is, $p\ val. = \frac{1}{N}\sum_{n=1}^{N} I_{\{W_n > W\}}$.

Regarding cointegration test, this is nonstandard for the presence of unit-root, nonlinearities and parameters unidentified under the null. According to Caner and Hansen (2001), dynamics of z_t can be re-written as

$$\Delta z_t = \left(\rho_1 I_{\{s_{t-d} \leq c\}} + \rho_2 I_{\{s_{t-d} > c\}}\right) z_{t-1} + \epsilon_t, \qquad (8)$$

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[51] Test statistics are taken from Caner and Hansen (2001). Furthermore, they argue how the residual bootstrap technique outperforms first order asymptotic techniques. Accordingly, we do not look for first order asymptotic approximation of test statistics, but we rely on the residual bootstrap procedure.
[52] $W = T\ \frac{S_0 - S_1}{S_1}$, where S_0 is the sum of squared residuals under the null, and S_1 under the alternative ($\delta_1 \neq \delta_2$). Caner and Hansen (2001) show the Wald statistics is also a super-Wald statistics, being S_T the supremum of Wald statistics as a function of the threshold c.
[53] Caner and Hansen (2001) also specified a restricted alternative to this test, where $\tilde{\delta} = 1$. Given that, both the power and the size were very similar to the restricted case, we rely on the latter.

where $\rho_1 = \delta_1 - 1$ and $\rho_2 = \delta_2 - 1$. The null H_0: $\rho_1 = \rho_2 = 0$ is taken to be no-cointegration, whereas two alternative hypothesis are specified. When z_t is stationary and ergodic in both regimes (i.e., threshold cointegration) the alternative is H_1: $\rho_1 < 0$, $\rho_2 < 0$. On the other hand, when there is a unit root only in one regime (i.e., the cointegration process is interrupted), the alternative can be specified as H_2: ($\rho_1 < 0$, $\rho_2 = 0$) or ($\rho_1 = 0$, $\rho_2 < 0$). We propose the test statistic R for testing H_0 against H_1 and H_2, with

$$R = t_1^2 \, 1_{\{\hat{\rho}_1 < 0\}} + t_2^2 \, 1_{\{\hat{\rho}_2 < 0\}},$$

where $t_1$and $t_2$are obtained from OLS regression of first equation in (1) as the t statistics of $\hat{\rho}_1$ and $\hat{\rho}_2$. R focuses on the one-sided alternative $\rho_1 < 0$ or $\rho_2 < 0$, considering only negative values of $\hat{\rho}_1$ and $\hat{\rho}_2$. This test is useful to recognize whether the data-generating process follows H_1 or H_2. Indeed, when R is significant we fail to accept the null of no-cointegration. Examining the individual t statistics $-t_1$ and $-t_2$, we account for interrupted cointegration (i.e., H_2) whenever just one of the two statistics is significant. Otherwise, the system follows the standard threshold cointegration process. Critical values for t_1, t_2, and R are calculated under the null H_0, based on a Monte Carlo simulation:

v) Estimation of the parameters of model (1) and (2) under H_0: $\rho_1 = \rho_2 = 0$;

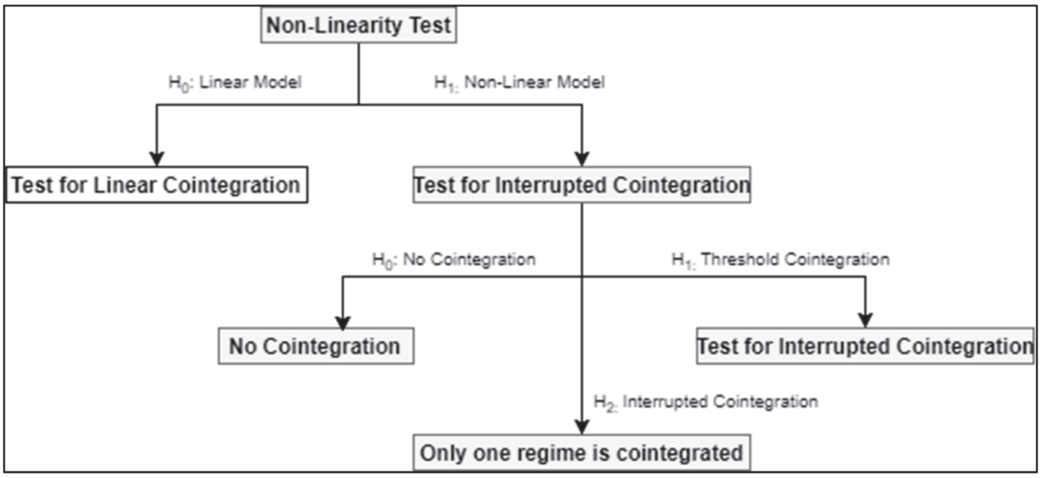
vi) Generation of $N = 20{,}000$ replications - with bootstrap re-sampling techniques – $\{x_t^*, y_t^*\}_n$, $t = 1, \dots, T$, $n = 1, \dots, N$, using the parameters estimated in v);

vii) Calculation of the test statistic $(t_1)_n, (t_2)_n, R_n$ for each series $\{x_t^*, y_t^*\}_n$;

viii) Calculation of the q% critical value of t_1, t_2, R from the q-quantile of the distributions of the test statistics $\{(t_1)_n, (t_2)_n, R_n, n = 1, \dots, N\}$. P-values can be calculated with the frequency of test statistic $(t_1)_n, (t_2)_n, R_n$ which are larger than given values of t_1, t_2, R.

Figure 10 illustrates and resumes methodological steps we followed. Interestingly, the linearity test is performed before the cointegration test, in contrast with traditional tests developed in Balke and Fomby, (1997), Hansen and Seo (2002), and Lo and Zivot (2001). Therefore, the discussed procedure does not depend on the results of standard linear-cointegration tests like the two-steps Engle-Granger test.

Figure 10 – Methodological Steps



Source: Authors' elaboration

3.5. Data

According to the applied methodology outlined in Section 4 and the literature review in Section 3, a large number of weekly time series-variables has been used for this study, with Italian corn and dairy feed prices constituting the backbone of our model (see the "In-Model Variables" in Table 25 and Figure 11). Three main categories are outlined in Table 25, trying to cover what literature considers as the primary drivers for price cycles: financial variables, energy-related prices, and agricultural commodities prices. The first group may help out in understanding whether concerns about agricultural financialization are significant for our specific case. As suggested in many empirical works (see among others Gilbert, 2010; Gilbert and Pfuderer, 2014; Mensi et al., 2014; Tang and Xiong, 2012), both the Dow Jones and S&P are the most tracked future indices[54]. Likewise, we considered the volume of open interests relative to corn and soybean futures on the U.S. market as a proxy for financial activity on certain agricultural commodities[55]. Futures[56] prices of corn and soybean were also investigated on both the U.S. and European markets. Within the energy-related prices, several proxies for biofuels price has been considered, both in the U.S. and the EU markets, accounting also for related futures. Brent and West Texas Intermediate crude oil prices were taken as standard oil prices. Both these two categories may well represent 'non-fundamentals' drivers, and as we already argued in Section 3, they can be interrelated since financialization may have strengthened the linkage food-fuel. Accounting for futures and spot prices for U.S. agricultural commodities help out in understanding whether the cereal market liberalization in the EU turns international prices significant for our VPT

[54] As described in Tang and Xiong (2012), the S&P index weights each commodity on its world production, whereas the DJ-index accounts for the amount of trading activity. For both, the energy sector represents a much greater weight than others.
[55] We relied upon the CBOT instead of the MATIF due to the negligible activity of the latter.
[56] Given the price discovery function of futures, they were also used as proxies for OTC prices when the latter were not available.

analysis. We rely on the American market since is this among the largest producers of corn, soybean, and ethanol. The last two categories, on the other hand, can be seen as part of the market fundamentals' category. In-model variables refer to the price of feed and feed-corn in Italy, and both were retrieved from the most representative production areas: the Bologna Chamber of Commerce for the Italian feed-corn price series, and the Chamber of Commerce of Forlì-Cesena for the feed price. They are both located in Emilia-Romagna which is the most representative region regarding feed production and among major corn-producer zones in Italy. Further agricultural commodity prices[57] were considered, given specific relationship they may have with feed production. Soybean is the second most crucial raw input for producing cattle feed, while rapeseed meal is increasing its presence in the feed industry due to the rise of biodiesel production[58]. Foreign corn prices for France and Ukraine are investigated since they represent 40% of total imports in 2016.

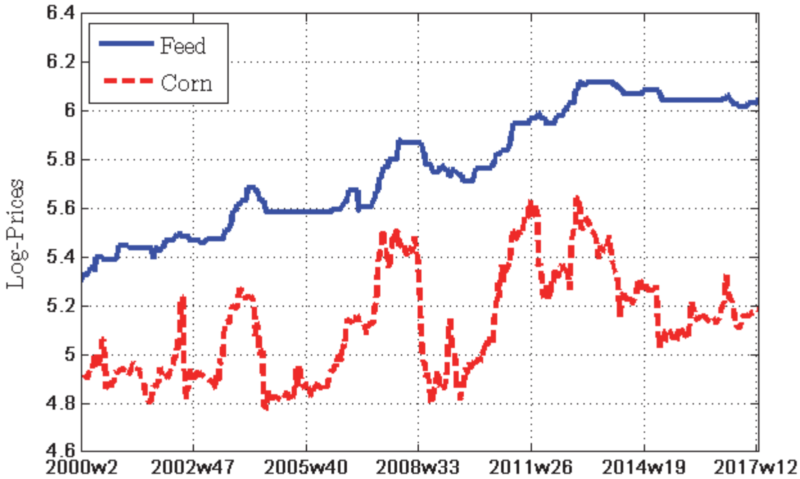
[57] We rely upon the German price for soybean meal given the latter is the largest producer in the EU. We also account for the Argentinean price of soybean being this one of the largest exporters.
[58] Due to the lack of data, rapeseed oil has been taken as a proxy of rapeseed meal. Likewise, soybean oil and and soybean meal futures as proxies for soybean meal.

Table 25 – Variables List and their Main Futures

	Description	Time	N. Obs	Source
In-Model Variables				
FC_IT	Italian Corn, Feed Use	2000-2; 2017-13	897	Chamber of Commerce of Bologna
Feed_IT	Compound Feed for Milk Cows	2000-2; 2017-13	897	Chamber of Commerce of Forlì-Cesena
Financial Variables				
CGSYSPT	S&P GSCI Price Index	2000-2; 2017-13	897	Thomson Reuters
DJCICOP	Dow Jones Commodity Index	2000-2; 2017-13	897	Thomson Reuters
CS.TOTF	Soyabeans Open Interest	2000-2; 2017-13	897	Thomson Reuters
SOY_CBOT	Soyabean Future CBOT	2000-2; 2017-13	897	Thomson Reuters
CC.TOTF	CORN OPEN INTEREST	2000-2; 2017-13	897	Thomson Reuters
PCOCS00	Corn Future MATIF	2000-2; 2017-13	897	Thomson Reuters
CORNYF1	Corn #.2 Yellow Future CBOT	2000-2; 2017-13	897	Thomson Reuters
Energy-Related Variables				
Oil_Brent	Europe Brent Spot Price FOB	2000-2; 2017-13	897	EIA
CZEC.01	Denatured Fuel Ethanol Future CBOT	2005-16; 2007-14	623	Thomson Reuters
HBISMER	Soybean Methyl Ester B100 (FOB Rotterdam)	2008-8; 2017-14	475	Thomson Reuters
BIOB100	Biodiesel B100 (FOB Midwst)	2013-3; 2017-13	190	Thomson Reuters
ETHANYH	Ethanol Spot Price (New York)	2006-26; 2017-12	559	Thomson Reuters
HBISMEA	Soybean Methyl Ester B100 (FOB Argentina)	2007-35; 2017-14	500	Thomson Reuters
HBIFAME	Fatty Acid Methyl Ester 100C (FOB Rotterdam)	2007-5;2017-12	528	Thomson Reuters
BIODMSD	Biodiesel B100 (FOB U.S.)	2013-2;2017-12	219	Thomson Reuters
WTI €	WTI Spot Price FOB	2000-2; 2017-13	897	EIA
Agricultural Commodity Prices				
SM_GER	Soyameal, German 44% prot., FOB Hamburg	2000-2; 2017-13	897	PublicLedger
RPNLEC1	Rapeseed Oil (FOB The Netherlands)	2001-50; 2017-13	796	Thomson Reuters
SOYBEAN	Soybeans No.1 Yellow Spot Price (U.S.)	2000-2; 2017-13	897	Thomson Reuters
UKR	Corn, Ukraine	2012-41; 2017-13	234	PublicLedger
FRA	Corn, France	2001-39; 2017-13	808	Thomson Reuters
SOYM_CBOT	Soybean Meal Future CBOT	2000-2; 2017-13	897	Thomson Reuters

Source: Authors' elaboration

Figure 11 – Italian Feed-Corn and Compound Dairy Cattle Feed, 2000-2017



Source: Authors' elaboration

According to econometrics methodology described in the previous section, the transition variable has to be stationary. Therefore, exception made for open-interests and the ECT value, they are all expressed in first differences and volatilities[59], being the latter another potential trigger.

3.6. Empirical Results

Series were transformed into their logarithmic forms, mitigating the fluctuations and increasing the likelihood of stationarity after first differencing (Hamilton, 1994). From an economic point of view, this also makes it possible to interpret results in percentage change terms (Ben-Kaabia and Gil, 2007). First methodological steps involve the test for unit-root[60] on both price series since only when both time series result to be I(1), cointegration framework can be applied. Results describe two non-stationary series. Thus, we proceed in testing for Granger causality (Granger, 1969), leading to the conclusion that corn granger causes feed price[61]. The two steps Engle-Granger Cointegration test (Engle and Granger, 1987) between feed (y_t) and corn (x_t) prices was then applied, failing to reject the null of no-cointegration[62].

Table 26[63] summarizes the results we obtained from the econometrics modeling. We report just those systems resulting in significant effects, that is for which there exists a

[59] Volatilities are estimated througha GARCH(1,1) process. See (Bollerslev, 1986) for more details.
[60] We employ three unit-root tests: the ADF (Augmented Dickey-Fuller) parametric test, the PP (Phillips and Perron, 1988) non-parametric test, and finally the KPSS (Kwiatkowski et al., 1992). For detailed results see the Appendix I.
[61] We chose the number of lags according to the BIC criterion, resulting in a p-value of 0.30 when causality goes from feed to corn, and 5.6621e-15 for the other way round, indicating a cost-push mechanism takes place.
[62] The p-value in the ADF test on the residuals of cointegrating equation equal to 0.48. For robustness check, even if we change the order of the variables we are not able to reject the null, concluding the two price series are not linearly cointegrated.
[63] For a more detailed summary of statistical results see the Appendix I

cointegrating framework. However, also for the non-significant variables, a discussion is provided.

Table 26 – Summary of Results for the (Interrupted) TAR Cointegration Models

Trigger Variable	Level	Threshold	% Obs. Regime 1	H-I 1	H-I 2
		In-Model Triggers			
Feed-Corn (IT)***	volatility d=1	3.5	84.2		14.3
ECT**	returns d=4	-3.6	15.0	16.9	
		Agricultural Commodities			
Soymeal (GER)**	returns d=1	-0.6	41.1		30.5
Soybean (U.S.)**	returns d=1	-0.2	42.9		32.2
		Energy-Related Triggers			
Rapeseed Oil (HOL)***	volatility d=1	3.5	81.5		11.9
WTI (U.S.)*	returns d=1	4.0	84.9		18.2
		Financial Triggers			
Corn_MATIF***	volatility d=1	3.2	84.9		12.6
Corn_MATIF **	returns d=4	-4.4	17.1	15.8	
DJ**	returns d=4	4.2	80.0		14.7
S&P **	returns d=4	5.7	82.2		15.6

The asterisks stand for statistical significance of the R test: *1%, **5%, ***10%

Source: Authors' elaboration

The variables listed in Table 26 above are those for which the Wald test of non-linearity rejects the null of linearity. Accordingly, when non-linearities arose, we test for cointegration in regime one and two. If we fail to reject the null of no-cointegration for both regimes, the system is a standard TAR model. Otherwise, that is if the null is not rejected just for one regime, we have an interrupted-cointegration system in which price series share a common stochastic trend only when the transition variable is above or below the estimated threshold value. The first regime refers to whenever the trigger variable is below the threshold value, where the second regime describes the opposite situation. Figure 12 illustrates the linkages between Italian corn and feed prices and the triggering variables.

The in-model transition variables

First differenced ECT depicts the well-known "rocket and feather" behavior described in the seminal paper of Peltzman (2000), so that only when margins are shrinking the feed price adjusts to the long-run equilibrium, re-establishing the steady state after around four months. This behavior is explained by the high concentration in the Italian feed industry, which uses its market power for maintaining high margins. Looking at the volatility of Italian corn prices, the system responds significantly only when the volatility is high, that is when risk and uncertainty in the market increase. Consistently with first results we mentioned, feed producers may want to be in-line with spot price movements as they expect price peaks, so they will be able to increase feed prices accordingly. Looking at Figure 11 the reader can easily argue how feed price increases are more pronounced than decreases.

On the other hand, nor feed price returns neither feed price volatility seem to be significant in triggering a response since the feed price strongly depends on feedstocks rather than its own-price. However, changes in Italian feed corn price do not trigger any significant response, what may be explained again by the exertion of market power by the feed industry. Additionally, one may charge that vertical integration and contracts buffer its impact.

Agricultural Commodities Linkages

Soy also enters into milk cow feed composition, and therefore we account for different prices relative to this commodity, like beans, and its by-products, meal. Soymeal price triggers significant responses when it increases. Here again, consistently with what we observed earlier, an increase in soymeal prices means feed industry’s margin is being squeezed, hence the price of feed re-adjusts to the long-run equilibrium. The same logic applies regarding the soybean price. Interestingly, no effects relative to both French and Ukrainian corn, a result consistent with Italian corn price. Indeed, French corn price appears to be in strictly co-movement with that of Italy, of course leading to similar results. On the other hand, the Ukrainian price series effects could be hidden by the length of the series (i.e., only 235 observations).

Energy-Related Commodity Prices

Interestingly, there is a substantial, significant effect of rapeseed oil price volatility. As reported by FEFAC (2017) the share of rapeseed meal has recently increased because of the surge of biodiesel production. Rapeseed meal is indeed the by-product of biodiesel production and is used as input for cattle feed production, substituting the soybean meal. However, given the low share of this type of input in the feed composition, its high significance is quite surprising. We argue that it shares the same volatility structure as Italian corn price, what turns it strongly significant. Indeed, no further biofuel price resulted in being significant. Concerning the crude oil price, despite the impact it may have on PT dynamics is not strongly significant (just at 10%). However, this enhances the idea that whenever it becomes more volatile, the feed price reacts. Probably the most exciting results hinge on the non-significant effect of some energy-related prices. The Brent crude oil price seems to have no impact on the corn-feed relationship, nor price changes neither price volatility. Furthermore, we found no significant impact of biofuels-related prices (FAME, SME, American Biodiesel price, and American Ethanol price). The latter can be due to the low ethanol production in the EU, and to the fact that biodiesel is mainly extracted from rapeseed rather than soybean. Furthermore, Europe does not extract ethanol from corn, and the volume produced is not that high to justify any significant effect.

Financial Variables

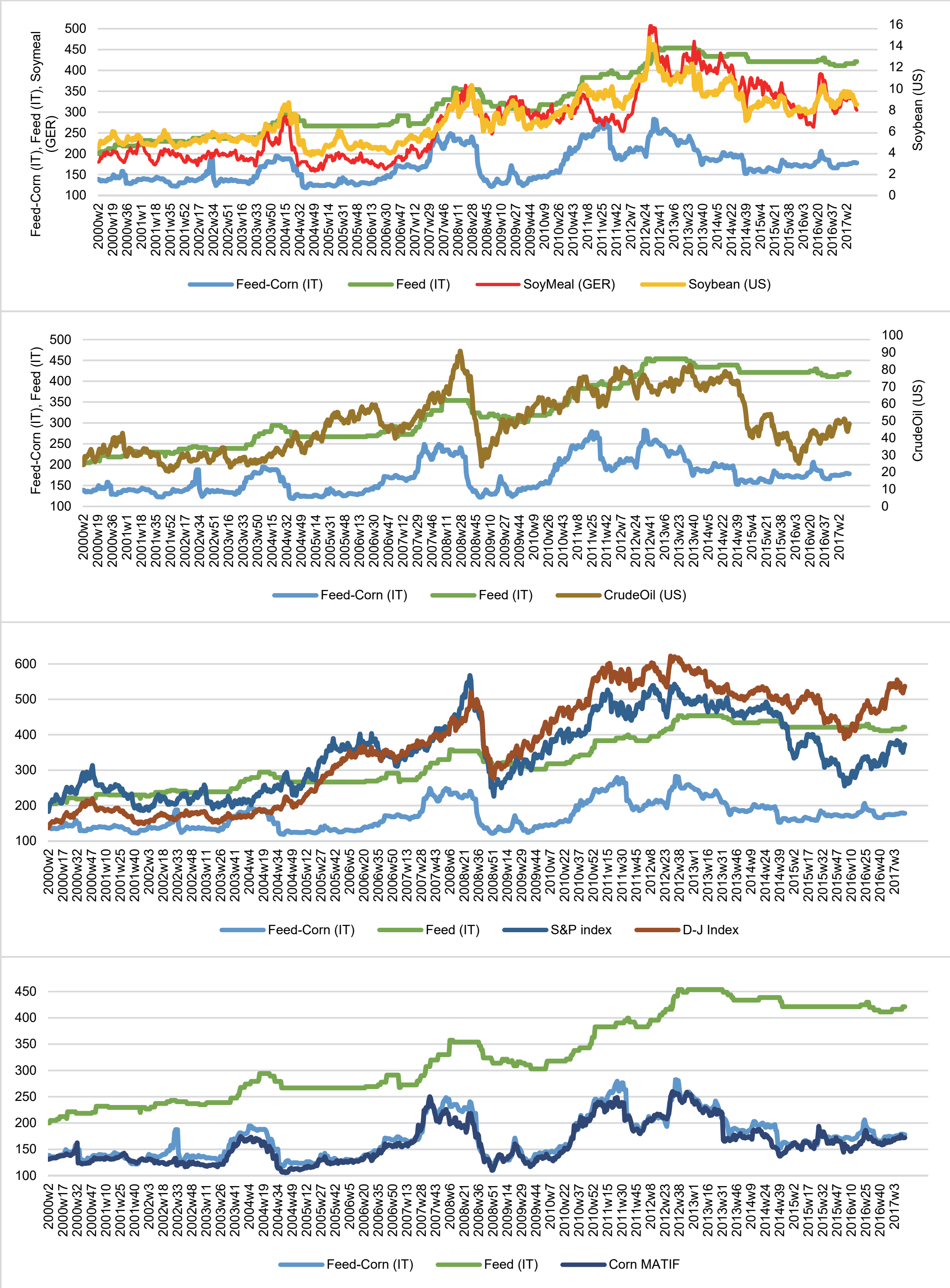
Both S&P and DJ commodity indexes result to be significant whenever positive changes in their quotations occur. These are two equity indexes, so their values are derived from a basket of commodities to which a particular weight is linked, based on future prices. However, energy prices represent a greater weight than other commodities for these two indexes (Gilbert and Pfuderer, 2014; Tang and Xiong, 2012). Bearing this in mind, we interpreted these results as the market expecting a general increase in agricultural commodities and energy prices, instead of a clear influence of financialization. According to this point of view, feed prices are also entitled to increase. In many empirical studies, futures were blamed for food price surge of 2006-2008. However, further to be derivative, and, hence, financial instruments, they also absolve an essential function of price discovery, that is allowing agents who have prior information to trade based on such information (Gilbert, 2010). Since they bear expectations of future spot prices, they are also significant for price formation on the spot market (Brümmer, Korn, Schlüßler, and Jamali Jaghdani, 2016). In light of the latter, MATIF corn futures reflects the reference price for corn in Europe, absolving their function of price discovery. Furthermore, a decrease in the European corn future triggers a response in the feed price. The central hypothesis behind this behavior is that a general reduction in the price of the primary input for feed production, must bring a reduction in the price of feed. A different behavior may entail severe disruption along the supply chain, with farmers inclined to retain a major part of maize for directly feeding their livestock. The impact of many financial variables turned to be nonsignificant. We found that the volume of open interests for corn and soybean futures on the CBOT has no effect on PT between Italian corn and feed prices. The first intuition behind the usage of this transition variables is that larger feed producers, also operating in foreign markets, may make use of American futures for risk management. Likewise, futures on soybean and its by-products, as well as for corn, resulting in a no-significant effect.

3.7. Conclusions

Price transmission analysis always featured the assumption of a continuous cointegrating relationship between prices, when it exists. However, due to high-volatile markets and structural changes, such relationship may be interrupted, showing periods of co-movements and others of random walk. In order to overcome such methodological drawback, we applied a brand new methodology that allows for a particular case of threshold cointegration in which one of the two regimes is non-stationary, in other words not cointegrated. According to an extensive literature review, we tried to account for all those drivers that may have a significant impact on the relationship between the Italian corn and dairy cattle-feed prices. Hence, we ended up distinguishing between fundamentals-related and non-fundamentals triggers. We find weak evidence of non-fundamentals drivers on VPT dynamics for the

Italian feed industry, agreeing with Wright (2011, p. 1) view on grain's economics, that is market fundamentals "[…] seems to be as relevant for our understanding of these markets as it was decades ago". Indeed, energy-related prices showed very weak linkages with the feed system. Biofuels play no significant role in shaping the corn-feed relationship, according to Serra and Zilberman (2013), Serra (2013), and López Cabrera and Schulz (2016). Concerning ethanol, Europe (and hence, Italy) does not represent a significant producer, since maize is almost entirely employed in the feed and food industries. Biodiesel, mainly produced from rapeseed instead of soybeans, neither represents a significant trigger for the investigated agrifood system. The crude American oil price has a quasi-non-significant effect, whereas the European Brent price plays no role, agreeing with conclusions drawn by Cooke and Robles (2009), Myers et al. (2014), Saghaian (2010), Zhang et al. (2010), Reboredo (2012), and Zhang et al. (2009). Financial variables neither seem to be of some importance for VPT dynamics, supporting general results as claimed by Gilbert and Pfuderer (2014), and Irwin (2013). We found significant effects from the two commodity indexes, but we argued such effects be mainly due to the general market expectations on commodity prices rather than a significant inference of financialization. However, such results agree with those scholars charging the financialization strengthened the relationship between food and fuel such as Tang and Xiong (2012), and Gilbert (2010). According to Ganneval (2016), high-volatility regimes trigger significant responses. Finally, if on the one hand, we have behaviors matching with market power exertion by feed processors, on the other hand, it seems that when European corn futures decrease, feed price follows the decreasing trend. Maize suppliers are, in many cases, the final users of feed. The feed industry may want to prevent disruptions along the supply chain. However, this work is far from being conclusive. Structural changes should be investigated to deepen the understanding whether certain drivers' impact changed from pre to post-crisis. Furthermore, a standard stock variable should be included as a transition variables as one of the most prominent market fundamentals. Regarding the econometric technique, more should be done for improving its potential. Firstly, the addition of one more threshold would provide more insights about the impact of the triggers. Second, short-run dynamics, i.e. Impulse Response Functions (IRFs) should be investigated in order to understand how the market behaves in the short-term according to the trigger variable deployed. Last but not least, generalization of the modeling approach to a multivariate system will definetely offer a more interesting methodology for the study of PT dynamics.

Figure 12 – Price Linkages among Italian Corn, Italian Feed, and Trigger Prices



Source: Authors' elaboration

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Conclusions

This Thesis faces pice transmission analysis under both the methodological and the economic angles, especially regarding the policy implications and the organization of the supply chain. The dairy market is a hot topic currently debated in the whole EU for its political and economic relevance. For each chapter a specific issue concerning the Italian milk sector is investigated, with the final aim of unveiling the complexity of this commodity market, offering a thorough analysis of vertical price transmission process in the Italian fluid milk supply chain.

We first looked at the retail level since it is often blamed for exerting market power, which is a primary source of market inefficiency. Moreover, we investigate whether transmission dynamics change according to product differentiation strategies, and we confronted the conventional fluid milk with its organic counterpart using retail-scanner price series over a period of more than fifteen years. We found significant differences, both in the short and the long-run, according to different product's characteristics and, hence, different retail structures.

EU agricultural policy has often been at the core of many institutional and societal debates and decisions, given its role in supporting and protecting European agricultural markets. The dairy sector has probably been among the most intervened sector in the CAP history, and the 2000s reforms, aimed at the overall liberalization of European agricultural markets, marked an essential structural change for the milk industry all over the EU. Providing insights about CAP's impacts on price transmission in Italy is of high importance for policymakers and stakeholders. On the one hand, results indicate a low-intervening policy brings a smoother transmission mechanism, eliminating asymmetric dynamics. On the other hand, responses to shocks are sluggisher after the liberalization, a consequence of the increase in volatility, and, hence, in uncertainty.

Finally, because of the world commodity price surge of 2007, exogenous variables such as energy prices and the financialization of agriculture, have been hypothetically interfering with market fundamentals in characterizing agricultural commodity price cycles. Given the significant interdependence between maize with the milk production, and the terrific price surge cereals experienced, we disentangled if and how exogenous elements impacted the transmission mechanism between corn and feed prices in the Italian scenario. Results point to the yet robust primary role of supply and demand in explaining price dynamics in the very first levels of the Italian milk chain.

Investigating a given agricultural commodity market is not an easy task, given the complex system of interrelations this has with many other market's elements, from strategic behaviors, and the related organization of the chain, to implications of *ad hoc* policies. The analysis of price transmission mechanisms, however, has the power to uncover many characteristics of the agricultural commodity one accounts for, since prices are the first linkage among all the actor of the market, both horizontally and vertically. Prices bear all the information available in the market, and the nature, magnitude, and speed of their transmission along the chain give a good measure of market integration, spotting inefficiencies and bottlenecks, hence representing an excellent instrument for policymakers, providing insights on where to intervene.

Despite the applied methodology relies upon a large number of sound empirical works, representing a cemented approach for price transmission analysis, different models – structural models, which hang on a different econometric approach, and further non-linear non-structural models - may entail different results. Indeed, a description of the advantages but also the drawbacks of non-structural time series econometric models have been provided. Furthermore, price data availability is still an issue for researchers, and of course the closer the dataset to the real prices, the more the results are robust and reliable, and they give a better picture of the real world.

Each section deals with different issues of the economic calendar and conclusions have been – by definition – drawn upon subjective economic hypothesis, although they strongly hinge on real economic facts. This is just a ground floor on which building up more reliable and robust analysis and approaches, to actively participate to institutional debates that eventually bring new policy and reforms, bridging the academic environment with the real-world state of play.

Appendix I

Table 27 – Unit Root Tests for Each Transition Variable

Variable	Measure	ADF	PP	KPSS	Variable	Measure	ADF	PP	KPSS
ETC	levels	0.020	0.019	<0.01	Soy OpenInterest	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
ETC	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Soybeanoil OpenInterest	h_t, d=1	0.518	0.517	<0.01
ETC	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Soybeanoil OpenInterest	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
WTI	h_t, d=1	0.247	0.241	<0.01	Soybeanoil OpenInterest	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
WTI	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Corn_futures	h_t, d=1	0.109	0.105	<0.01
WTI	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Corn_futures	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Soymeal (DE)	h_t, d=1	0.247	0.250	<0.01	Corn_futures	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Soymeal (DE)	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Soy_oil_Arg	h_t, d=1	0.636	0.638	<0.01
Soymeal (DE)	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Soy_oil_Arg	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Corn (IT)	h_t, d=1	<0.001	<0.001	<0.01	Soy_oil_Arg	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Corn (IT)	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Soy (US)	h_t, d=1	0.172	0.164	<0.01
Corn (IT)	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Soy (US)	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Feed (IT)	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Soy (US)	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Feed (IT)	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Soy_CBOT	h_t, d=1	0.035	0.038	<0.01
ECBOT-ETHANOL	h_t, d=1	0.053	0.042	<0.01	Soy_CBOT	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
ECBOT-ETHANOL	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Soy_CBOT	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
ECBOT-ETHANOL	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	SoyMeal_CBOT	h_t, d=1	0.118	0.109	<0.01
Soy_Methyl_Rott	h_t, d=1	0.588	0.572	<0.01	SoyMeal_CBOT	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Soy_Methyl_Rott	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	SoyMeal_CBOT	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Soy_Methyl_Rott	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	SoyOil_CBOT	h_t, d=1	0.350	0.350	<0.01
Ethanol_NY	h_t, d=1	0.206	0.209	<0.01	SoyOil_CBOT	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Ethanol_NY	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	SoyOil_CBOT	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Ethanol_NY	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	WTI	h_t, d=1	0.281	0.286	<0.01
RapeseedOil	h_t, d=1	0.097	0.101	<0.01	WTI	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
RapeseedOil	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	WTI	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
RapeseedOil	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Corn OpenInterest	h_t, d=1	0.445	0.413	>0.1

Table 27 – Cont'd

Soy_Methyl_Arg	h_t, d=1	0.482	0.459	<0.01	Corn OpenInterest	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
Soy_Methyl_Arg	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Corn OpenInterest	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
Soy_Methyl_Arg	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	MATIF-CORN	h_t, d=1	<0.001	<0.001	<0.01
FAME	h_t, d=1	0.418	0.421	<0.01	MATIF-CORN	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
FAME	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	MATIF-CORN	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
FAME	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Corn (FR)	h_t, d=1	<0.001	<0.001	>0.1
S&P_GSCI	h_t, d=1	0.299	0.294	<0.01	Corn (FR)	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
S&P_GSCI	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Corn (FR)	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
S&P_GSCI	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	Mais-France-2	h_t, d=1	0.002	<0.001	<0.01
DJ	h_t, d=1	0.307	0.306	<0.01	Mais-France-2	Δx_{t-i}, d=1	<0.001	<0.001	>0.1
DJ	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	Mais-France-2	Δx_{t-i}, d=4	<0.001	<0.001	>0.1
DJ	Δx_{t-i}, d=4	<0.001	<0.001	>0.1	STD	h_t, d=1	0.536	0.536	<0.01
Soy OpenInterest	h_t, d=1	0.430	0.384	<0.01	STD	Δx_{t-i}, d=1	0.553	0.553	>0.1
Soy OpenInterest	Δx_{t-i}, d=1	<0.001	<0.001	>0.1	STD	Δx_{t-i}, d=4	<0.001	<0.001	>0.1

$\boldsymbol{h_t}$ is volatility, estimated as a GARCH(1,1) process; $\Delta \boldsymbol{x_{t-1}} = \boldsymbol{x_t} - \boldsymbol{x_{t-1}}$; d is the delay parameter

Source: Authors' elaboration

Table 28 – Garch (1,1) Results for Each Transition Variables

Transition Variable	Ω	α	β	α + β[1]	Transition Variable	Ω	α	β	α + β[1]
WTI	0.00*	0.094*	0.895*	0.990	Soy OI	0.001*	0.062	0	0.062
Soymeal (DE)	0.00*	0.077	0.817*	0.895	Soyoil OI	0.00*	0.027	0.945*	0.972
Corn (IT)	0.00*	0.293*	0.706	1.000	CBOT Corn (US)	0.00*	0.125	0.726*	0.851
Feed (IT)	0.00*	0.012	0.075*	0.088	Soy_oil_Arg	0.00*	0.033	0.937*	0.971
ECBOT-ETHANOL	0.00*	0.189	0.642*	0.832	Soy (US)	0.00*	0.109*	0.828*	0.937
Soy_Methyl_Rott	0.00*	0.069*	0.889*	0.959	Soy_CBOT	0.00*	0.154*	0.798*	0.953
Ethanol (US)	0.00*	0.101*	0.841*	0.943	SoyMeal_CBOT	0.00*	0.125*	0.813*	0.939
RapeseedOil	0.00*	0.128*	0.812*	0.941	SoyOil_CBOT	0.00*	0.074*	0.880*	0.955
Soy_Methyl_Arg	0.00*	0.077*	0.896*	0.973	WTI	0.00*	0.079*	0.896*	0.975
FAME	0.00*	0.033*	0.960*	0.994	Corn OI	0.00*	0.048	0.201*	0.249
S&P	0.00*	0.078*	0.884*	0.963	MATIF-CORN	0.00*	0.323*	0	0.323
DJ	0.00*	0.068*	0.903*	0.972	STD	0.122	0.981	0.000	0.981
Corn (FR)	0.00*	0.338*	0.146*	0.485					

[1] Whenever $\alpha + \beta < 1$ the process is weak-stationary. If $\alpha + \beta = 1$ the process is an IGARCH, and the unconditional variance does not exist. However, the IGARCH process can be strongly-stationary even though it is not weakly-stationary. See Nelson (1991) for details.
For a thorough understanding of GARCH estimates see Bollerslev (1986).

Source: Authors' elaboration

Table 29 – Model Results for Each Significant Transition Variable

Transition Variable	Measure	μ[1]	β[1]	Wald (p-value)	c	% Obs. 1st Regime	HL[2] δ 1	HL[2] δ 2	R (p-value)	T1 (p-value)	T2 (p-value)
RapeseedOil	h_t, d=1	2.488 (17.953)	0.646 (23.992)	0.001	3.545	82		11.941 (0.944)	0.002	0.739	**0.001**
MATIF	h_t, d=1	1.867 (13.613)	0.761 (28.439)	0.002	3.188	85		12.57 (0.946)	0.002	0.754	**0.001**
Corn(IT)	h_t, d=1	1.867 (13.613)	0.761 (28.439)	0.013	3.486	84		14.343 (0.953)	0.011	0.626	**0.004**
MATIF	Δx_{t-i}, d=4	1.889 (13.613)	0.757 (28.439)	0.017	-4.352	17	15.799 (0.957)		0.015	**0.005**	0.709
DJ	Δx_{t-i}, d=4	1.889 (13.613)	0.757 (28.439)	0.013	4.217	79		14.727 (0.954)	0.015	0.667	**0.005**
S&P	Δx_{t-i}, d=4	1.889 (13.613)	0.757 (28.439)	0.028	5.719	82		15.6 (0.957)	0.026	0.64	**0.007**
Soymeal (DE)	Δx_{t-i}, d=1	1.867 (13.613)	0.761 (28.439)	0.017	-0.635	41		30.503 (0.978)	0.032	0.959	**0.011**
ETC	Δx_{t-i}, d=4	1.889 (13.613)	0.757 (28.439)	0.038	-3.584	15	16.856 (0.96)		0.034	**0.016**	0.625
Soybean (US)	Δx_{t-i}, d=1	1.867 (13.613)	0.761 (28.439)	0.033	-0.228	43		32.172 (0.979)	0.047	0.938	**0.018**
WTI	Δx_{t-i}, d=1	1.867 (13.613)	0.761 (28.439)	0.088	4.026	85		18.155 (0.963)	0.077	0.58	**0.033**

$\boldsymbol{h_t}$ is volatility, estimated as a GARCH(1,1) process; $\Delta \boldsymbol{x_{t-1}} = \boldsymbol{x_t} - \boldsymbol{x_{t-1}}$; d is the delay parameter.
[1]Values in brackets are t-statistics.
[2]Values in brackets are the estimated coefficients ρ.

Source: Authors' elaboration