On Policy Interventions and Vertical Price Transmission: the Italian Milk Supply Chain Case

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Abstract

During the last two decades, the EU dairy sector has been interested by considerable changes and two policy reforms, the Fischler Reform and the Common Market Organization Reform, pushing toward economic liberalization. These changes affected the EU supply chains at different levels, altering the mechanisms of vertical price transmission. Against this background, we apply error correction models to assess how price signals are passed through, before and after the Italian milk supply chain reforms. In particular, we study the degree of price transmission asymmetries and conclude that market sluggishness has increased in the post-reform period, but the asymmetric dynamics are less evident. Reflections on future research needs are discussed.

Keywords: asymmetries, CAP reform, dairy sector, error correction model, Fischler reform, structural break

JEL Code: L16, Q13, Q18

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Introduction

The Italian dairy industry, especially the fluid milk sector, has experienced relevant changes during the last decades. Searching for liberalization is not new in the EU (*cfr*. the McSharry Reform in 1992) (Swinbank, 1993; Coleman and Tangermann, 1999) and the 2003 Reform, mainly due to the abolishment of the quota regime, enhanced the farm-gate competitiveness and developed a more market-driven dairy industry (Henke et al., 2018). It has also stiffened the competition among intra- and extra-EU producers and altered the distribution of margins along the supply chain. These changes are well captured by price signals, whose dynamics (pre- and post- Reform) may help infer the functioning of the supply chain.

The analysis of vertical price transmission (VPT) has a long tradition in agricultural economics (*cfr*. Lloyd, 2017, for a recent review). Prices are signals generated by economic transactions and convey the (explicit and sunk) information available on the market. The degree and the speed of price transmission proxy the degree of integration of the supply chain and inform on market efficiency, a key feature to plan strategic and structural market interventions (Goodwin and Holt, 1999; Abdulai, 2002; Serra and Goodwin, 2003).

The literature on price transmission is vast, but the majority of studies has focused on spatial/horizontal price transmission (Brester and Wohlgenant, 1997; Awokuse, 2007; Cioffi, Santeramo, and Vitale, 2011; Santeramo and Cioffi, 2012; Chen and Saghaian, 2016; Esposti and Listorti, 2018; Santeramo and Di Gioia, 2018; Durborow et al., 2020; Hatzenbuehler, Du, and Painter, 2020), with a few of them related to the dairy sector (Fabiosa et al., 2005; Kempen et al., 2011; Bergmann, O'Connor, and Thümmel, 2015; Hillen and von Cramon-Taubadel, 2019). The applications of the VPT analysis to the food supply chains are also numerous (e.g., Ben-Kaabia and Gil, 2007; Boetel and Liu, 2010; Li and Sexton, 2013; Ahn and Lee, 2015; Rezitis, 2017; Antonioli et al., 2019), but only a few studies have analyzed the impacts of the

CAP reforms (e.g., Cacchiarelli et al., 2016; Rezitis and Pachis, 2018), especially concerning the milk supply chain[1].

We explore how the Common Agricultural Policy (CAP) Reform of 2003, intended to liberalize the EU dairy markets, has influenced the VPT dynamics in the Italian context. Given the perishable nature of fluid milk, a constrained supply side due to the quota regime, and quasi-fixed production processes due to contracts, the Italian fluid milk market has suffered from an unbalanced bargaining power exerted by downstream agents. The Italian milk market's liberalization has accelerated structural changes, with mergers and acquisitions at farmer and processor levels (OMPZ, 2019) that may explain observed changes in bargaining power and price transmission dynamics. The unbalanced market power, which favour the retailers, the rigidity of the supply, as implied by policy measures, and the low storability of the fresh fluid milk, are likely to have favoured asymmetric price transmission dynamics[2]. The Reform has liberalized the market and likely favored a whole, symmetric, fast transmission of market signals.

We investigate the dynamics of VPT in the Italian milk market, using processors and consumers' prices for fresh fluid milk from January 2000 to August 2016. We assess the nature of price transmission (i.e., cost-push, demand-pull, or feedback system), the degree the symmetry, and the speed for shocks to be passed-through.

[1] Albeit the scant literature regarding policy impacts, price transmission along the milk supply chain is widely investigated: see among others Lloyd et al., 2009; Richards, Allender, and Hamilton, 2012; Loy, Weiss, and Glauben, 2016.

[2] According to OMPZ (2009, 2019), during the period 2000-2016, the CR4 for dairy processors increased from approximately 12% to 15%, while food retailers featured an increase in CR4 from 38% to 45% (AGCM, 2013; MedioBanca, 2020), maintaining market power. Finally, the farm population shrank by almost 70% (approximately 29,000 units in 2016), with small milk farms exiting the market or selling their activity to bigger units (the number of heads per farm doubled since 2000, being approximately 45 heads per farm in 2016).

The Dairy Sector: policy and economic facts

The dairy sector has been liberalized through the CAP reform of 2003, referred to as Fischler Reform. As stated by the European Court of Auditors (2009, p. 13), "[…] *the 2003 reform initiated the liberalization of the milk sector by reducing price support and creating direct income support*". After the Reform, the milk target price has been removed, the intervention prices for dairy products have been lowered, and the national milk quotas have been dismantled (Gohin and Latruffe, 2006; DG-AGRI, 2010). A few years later, in 2007, the new Single Common Market Organization had a significant impact on the European milk market, setting to zero the export subsidies for milk (exception made for exports made from January to November 2009), besides new intervention prices for butter and skimmed milk powder (DG-AGRI, 2011; Meijerink and Achterbosch, 2013). On March 1st of 2015, the quota system, introduced in 1984[3] to regulate the supply surplus and sustain farm-level prices in the EU[4], has been dismantled[5] (Giles, 2015).

The Total Support Estimate (TSE), an OECD indicator that combines all agriculture-related public expenses, and the Producer Support Estimate (PSE) [6], which proxies the direct support paid to agricultural producers, have declined tremendously in 2007 (Figure 1). These dynamics

[3] For more details on the aforementioned policy interventions, please see Council Regulation (EC) No. 1234/2007, Council Regulation (EC) No 1255/1999, EC Regulations 856/84 and 857/84.

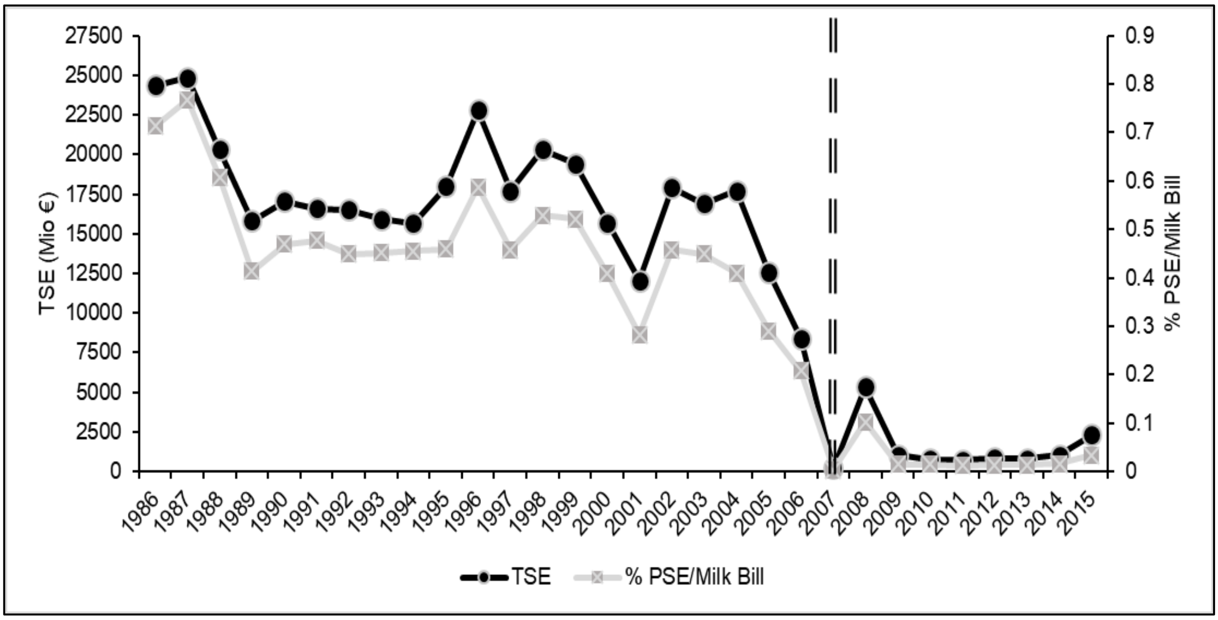
[4] The so-called quota-rent: the amount of rent generated from a restriction on supply (Tonini and Domínguez, 2009).

[5] The abolition of the quota has been preceded, in 2008, by phasing-out measures: quotas have been increased by 1-2% for five consecutive years (exception made for Italy, for which there has been a 5% increase in the 2009-10 campaign) (European Commission, 2010).

[6] See the OECD's PSE Manual at http://www.oecd.org/tad/agricultural-policies/psemanual.htm.

suggest that the CAP has moved to a far less-intervening policy framework: since 2007, the EU dairy sector has received almost no support.

Figure 1: TSE and PSE in the EU dairy sector



Source: Authors elaboration on OECD.

The Italian dairy industry's value is around 15 billion euros (11.4% of the food industry). The Italian industry embraces more than two thousand dairy firms and thirty thousand employees and accounts for more than 7% of the total EU-28 production, although remaining a net importer of dairies. Raw milk is sold as fluid milk (18%) or devoted to cheese production (70%). Most of the milk exchanged on the Italian market between farmers and industrial processors is regulated through contracts or marketed by cooperatives (OMPZ, 2019). The Italian Institute of Statistics (ISTAT) provides a nationally representative milk price index that averages the prices received by the industrial processors, a few big companies that collect a large (and increasing) volume of the marketed milk (OMPZ, 2019).

Literature Review

Gardner' seminal paper (1975) on price transmission warns on the importance of studying price dynamics along the supply chain and has stimulated a large volume of empirical studies, recently reviewed by Lloyd (2017). The VPT may be asymmetric due to the strategic interactions across stakeholders operating at different supply chain stages (Serra and Goodwin, 2003). Asymmetries may signal unfair welfare distribution (European Commission, 2009; OECD, 2015), but tend to be more common than expected (Peltzman, 2000; Bakucs, Falkowski, and Ferto, 2014). Meyer and von Cramon-Taubadel (2004) point to market power as a significant source of asymmetric price adjustments (see Sckokai, Soregaroli, and Moro (2013) for a market-power study referring to the Italian dairy sector). Peltzman (2000) and Serra and Goodwin (2003) found evidence of asymmetries for non-perishable products and symmetric VPT for high-perishable milk products. Kim and Ward (2013) conclude that VPT in fruit and vegetable commodities is asymmetric, with decreases in wholesale prices passed through more quickly to retailers than price increases. Similarly, Ahn and Lee's (2015) found that highly-perishable fruits are characterized by negative asymmetries, while the opposite occurs for low-perishable ones. Santeramo (2015) supports Kim's and Ward (2013) conclusions for EU vegetable markets: wholesalers price decreases impact more on retail than price increments. Santeramo and von Cramon-Taubadel (2016) conclude that symmetric PT is more common for highly perishable products. Empirical applications aimed at understanding how asymmetries react to policy interventions are rare (Vavra and Goodwin, 2005). Kinnucan and Forker (1987) conclude that government support for producer prices (e.g., floor prices) may lead to APT. Santeramo and Cioffi (2012) and Cioffi et al. (2011) concluded that the system of quotas and tariffs in the fruit and vegetable sector alter price dynamics for the imported products. Similarly, Lee and Gómez (2013) argue that the abolition of the coffee quota system altered VPT mechanisms and made retail prices more responsive to international price changes. Cacchiarelli et al. (2016) investigated the impacts of the mid-term CAP reform and concluded on the symmetry (asymmetry) of farm-wholesaler (wholesaler-retailer) prices. Finally, consumers' reaction to positive and negative price movements may induce asymmetric VPT (McCorriston, 2015), with evidences also found for the dairy sector (Biden et al., 2020).

To sum up, price transmission is informative on how policy interventions may alter market fundamentals, representing a useful measure for understanding the investigated supply chain's functioning. We contribute to the literature by investigating how the EU policy reforms alter price transmission along the supply chain.

Methodology

Following the literature on VPT, we relate consumer prices (P_c) with processor prices (P_p), allowing for asymmetric dynamics. Autocorrelations, unit-roots, and long-run price relationships (Engle and Granger, 1987) are accounted through an Error Correction Model (ECM):

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

where $P_{c,t-1} - \alpha_0 - \beta_1 P_{p,t-1} = ECT_{t-1}$ represents the Error Correction Term. Equation (1) assumes producer price to be exogenous and the cost-push mechanism to lead the PT dynamics (*cfr.* Ben-Kaabia and Gil, 2007; Santeramo and von Cramon-Taubadel, 2016). The assumption is empirically tested, as described in Section 6.

The parameters in equation (1), assumed to be constant over time, are tested against structural breaks. We allow for regime-dependent linear models (Asymmetric Error Correction Model, AECM) with positive and negative deviations in the Error Correction Term (ECT) (Granger and Lee, 1989):

$$\Delta P_{c,t} = \alpha_0 + \alpha_1 ECT_{t-1}^{+} + \alpha_2 ECT_{t-1}^{-} + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \Delta P_{p,t-i} + \varepsilon_t \quad (2)$$

Where $ECT_{t-1}^{+} = \begin{cases} ECT_{t-1} \; if \; ECT_{t-1} > 0 \\ 0 \; otherwise \end{cases}$, and $ECT_{t-1}^{-} = \begin{cases} ECT_{t-1} \; if \; ECT_{t-1} < 0 \\ 0 \; otherwise \end{cases}$.

We test the null of symmetry in the long-run ($H_0: a_1^{+} = a_2^{-}$) versus the alternative hypothesis of asymmetry. We also allow for asymmetries in the short-run (AECMSR) and test the null hypothesis of symmetry in the short-run ($H_0: \psi_1 = \psi_2$) against the alternative of asymmetry:

$$\Delta P_{c,t} = \alpha_0 + \alpha_1 ECT_{t-1}^{+} + \alpha_2 ECT_{t-1}^{-} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_1 \, \Delta P_{p,t-1}^{+} + \sum_{i=1}^{k-1} \psi_2 \, \Delta P_{p,t-1}^{-} + \mu_t \qquad (3)$$

Modeling structural breaks is crucial to avoid potentially biased results (Boetel and Liu, 2010; Lence, Moschini, and Santeramo, 2018; Liu, Chen, and Rabinowitz, 2019). Breaks in the long-run cointegration relationships, possibly due to the policy interventions, are tested through the Zivot and Andrews (1992) and modeled via an AECM with a structural break in the ECT:

$$\Delta P_{c,t} = \alpha_3 + (1 - D_t) \cdot (\alpha_1 ECT_{t-1}^{+} + \alpha_2 ECT_{t-1}^{-}) + D_t \cdot (\alpha_3 ECT_{t-1}^{+} + \alpha_4 ECT_{t-1}^{-}) + \sum_{i=1}^{k-1} \Gamma_j \, \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_j \, \Delta P_{p,t-i} + \mu_t \qquad (4)$$

where $D_t = \begin{cases} 1 & if \quad t \geq t_{bp} \\ 0 & if \quad t < t_{bp} \end{cases}$ is the Heaviside indicator function, and t_{bp} is the break-point. We disentangle price transmission dynamics preceding ($D_t = 0$) and following ($D_t = 1$) the policy change and test the null hypothesis of "no-changes".

Preliminary Analysis

In the post-Reform period, prices turned highly volatile[7]. Consumer prices increased by 28% (the volatility, σ_t, in the pre-, $\sigma_{Jan\ 00-Jul\ 07}$, and post-Reform periods[8], $\sigma_{Aug\ 07-Aug\ 16}$, were 0.33 and 0.42, respectively). The processor price's volatility rose by 83% ($\sigma_{Jan\ 00-Jul\ 07} = 0.37$, and $\sigma_{Aug\ 07-Aug\ 16} = 0.69$), suggesting that the retailers do not entirely pass to consumers the industrial processor price changes.

Out of 200 observations – 92 and 108 observations related to the pre- and post-Reform periods, respectively – 194 refer to price changes – 90 and 104 occurring in the pre- and post-break periods, respectively. The graphical analysis on price margins (Figure 2) shows that positive margin changes have been more frequent in the pre-break period, while they have been lower in magnitude (0.37€/l on average) when compared to the second period (0.49€/l on average). The margin increases (dotted white bars) are less persistent after 2007: the average period of occurrence is 2.9 months after the break, compared to 6.3 months before the break. Furthermore, more than 60% of the negative changes in margins occurred in the post-2007 period, characterized by a more symmetric evolution of margin changes.

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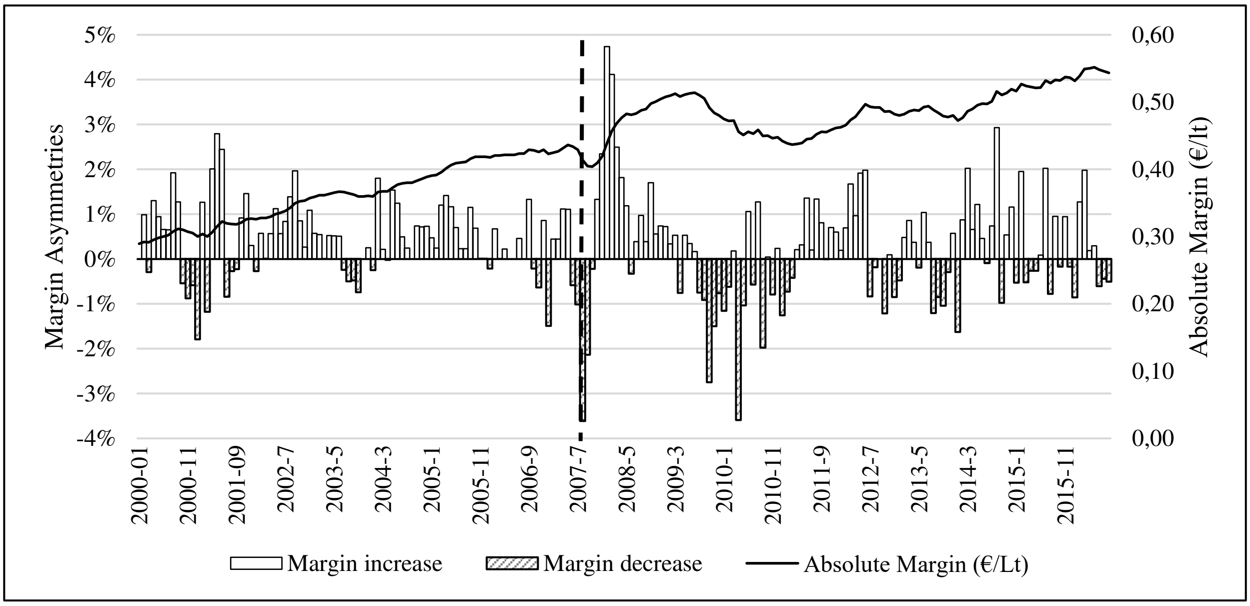
[7] Volatility has been calculated as: $\sigma_t = \sqrt{\frac{1}{m}\sum_{i=1}^{m} r_{t-i}^2}$, where $r_t = 100 \cdot (P_t - P_{t-1})/P_{t-1}$, m is the number of observations and the mean of returns $\bar{r}$ is assumed to be 0.

[8] Both the pre- and post-Reform periods refer to the structural break empirically detected: August 2007. See the following Section for more details.

Figure 2: Italian Milk Margin Trend



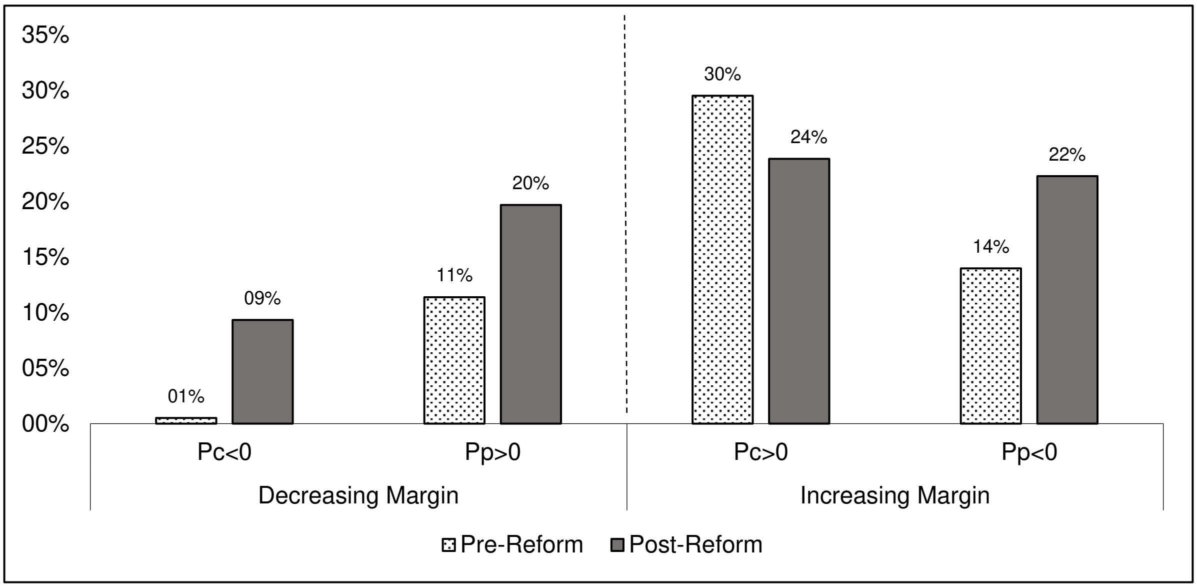
Notes: We define absolute margin as $M_t = P_{c,t} - P_{\mathrm{p,t}}$, and margin asymmetries as percentage changes $\Delta M_t = \left[\frac{(P_{c,t}-P_{\mathrm{p,t}})-(P_{c,t-1}-P_{\mathrm{p,t-1}})}{M_{t-1}}\right] \cdot 100$.

Source: Authors' elaboration

In the pre-reform period, positive margins are 2.8 times decreasing margins, whereas after the Reform took place, the ratio is 1.4; indeed, the periods of margin decrease doubled, from 23 to 44. The consumer price (Pc) upward movements shrunk by 17% after the Reform, whereas Pc reductions are 15 times more than the pre-reform scenario. Processor prices (Pp) increased about one-fourth regarding the positive price movements, with an increase of about 70% in price decreases.

In the post-Reform period (Figure 3), price changes are (generally) more frequent (55%) with respect to the pre-Reform period - and more evenly distributed, signaling a more balanced bargaining power along the chain.

Figure 3: Margin changes and price dynamics in pre- and post-Reform periods



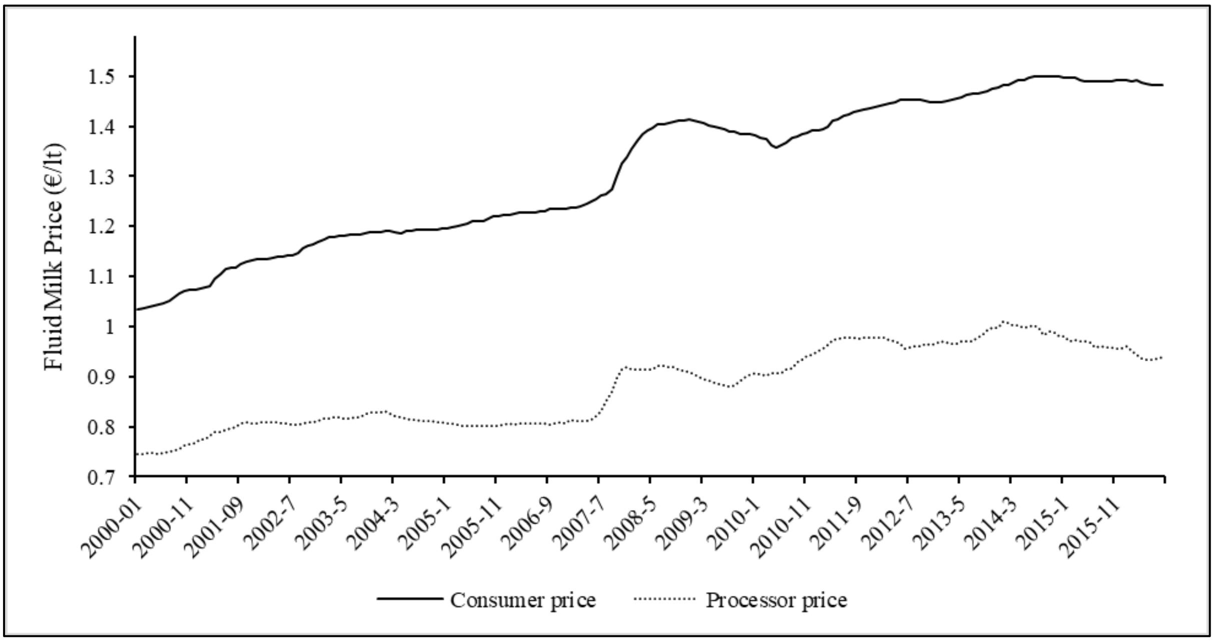
Source: Authors' elaboration

Data and Results

We use monthly data, from January 2000 to August 2016, of the price indexes (August 2010=100) of the fluid milk supply chain (Figure 2): the price paid by retailers to fluid milk processors (P_p) and the price applied by the retailers to consumers (P_c), have been collected from the Italian National Institute of Statistics (ISTAT)[9].

[9] Notably, given that industrial processors do not perform significant transformation to the fluid milk collected from farmers – pasteurization and packaging – our processor price is a valid proxy of the farm-gate price.

Figure 2: Consumer and Processor Prices of Fresh Fluid Milk in the Italian Market, 2000-16.



Note: For a more comprehensive graphical representation, we multiplied the two indexes by the absolute nominal price of fluid milk observed in August 2010, which is the base on which the two indexes rely upon. The absolute market price for both the supply chain level refers to the average price registered by one of the most representative food retailers in the Italian context at both the consumer and processor level (personal communication).
Source: Authors' elaboration.

The unit-root tests with constant (proxying price margins) suggest that the series are integrated of order one (I(1))[10]. The Zivot-Andrews test points to structural breaks in June 2007 (processor price) and September 2007 (consumer price)[11]: we use a middle point (August 2007) as a break-point in the cointegration relationships. The cointegration test fails to reject the existence of one cointegrating relationship (see Table 1). By normalizing on consumer price[12], we obtain $P_c = 0.206P_p + 0.149D_t + 3.576$, where D_t is a 0-1 dummy variable taking the value of 1

[10] We employed diverse unit-root tests: the GLS-ADF test (Elliott, Rothenberg, and Stock, 1996), the KPSS test (Kwiatkowski et al., 1992), and the PP test (Phillips and Perron, 1988), all clearly pointing to I(1) series (see Table A1 in the Appendix).

[11] See Table A3 in the Appendix for detailed results.

[12] Both the LM test for autocorrelation, and the LM-ARCH test for heteroskedasticity convey satisfactory results, pointing to a well-specified model. Results are available upon request.

after August 2007, when the structural break occurs[13]. The weak-exogeneity test allows concluding on the exogeneity of the processor price, signaling a cost-push mechanism, in line with most studies on PT in agricultural markets[14].

Table 1: Results of Cointegration Test with Structural Breaks

Rank	Trace	P-Value
0	27.832	0.033
1	8.224	0.266

Note: Constant restricted into the cointegrated space, together with the dummy variable D_t taking the value of 1 when $\text{t} \geq 2007:08$ (i.e., the structural break);
Number of lags included: 2 (based on the Schwartz Information Criteria)

Source: Authors' elaboration

The preliminary results on the ECM model (see equation (1)) suggest that a long-run relationship links prices (see Table 2, panel a)[15], where about 12% of a change in processor prices is transmitted to consumer prices. This suggests a sluggish adjustment of the system towards the equilibrium once a shock occurred.

The asymmetric short-run model (ASRM), specified as $\Delta P_{c,t} = \alpha_0 + \sum_{i=1}^{k-1} \Gamma_i \, \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_1 \, \Delta P_{p,t-1}^{+} + \sum_{i=1}^{k-1} \psi_2 \, \Delta P_{p,t-1}^{-} + \mu_t$ (Table A8, panel a.1), suggests that positive changes in processor prices influence consumer prices: when processor prices increase (and margins

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[13] We also investigate the presence of cointegration restricting a trend into the cointegrating space, both with and without a structural break, concluding on no cointegrating relationship. See Table A4 in the Appendix for further details.

[14] In order to reinforce the exogeneity assumption, Granger-Causality test was investigated, leading to the exogeneity of processor price. Moreover, the marginal model (see von Cramon-Taubadel, 1998) estimated for sake of model consistency point to exogenous processor price. Accordingly, we set P_c as the independent variable. See Table A2 and Table A5 in the Appendix.

[15] The ECT parameter is statistically significant. We computed the Error Correction Term from the linear cointegrating regression, $ECT_t = P_{c,t} - \alpha_0 - P_{p,t} - D_\text{t} = P_{c,t} - \alpha_0(-0.004) - P_{p,t}(0.994) - D_\text{t}(0.040)$.

squeeze), consumer prices tend to react. In general, changes in processor prices positively influence consumer prices, though the increase is not fully transmitted.

The dynamics are consistent with the AECM results (equation 2) (Table 2, panel b): when margin squeezes (negative ECT), consumer prices react toward the long-run equilibrium. The F-test on asymmetry suggests that price adjustments are asymmetric. In panel c, the model is estimated including asymmetric dynamics in both the ECT and the short-run (see equation (3)): results indicate that processor price increases and negative ECT (i.e., margins squeeze) influence consumer prices, consistently with results from panel b. All in all, we found asymmetries in both long and short-run price adjustments, as the F-test suggests in the last two rows of panel c, a result that is in line with the price behavior described by the literature: prices behave like "feathers" when margins stretch, and like "rockets" when they squeeze (Peltzman, 2000).

Table 2: Results from the Estimated Models: ECM, ASRM, AECM, AECMSR and AECMSB

	(a) ECM		**(b) AECM**		**(c) AECMSR**		**(d) AECMSB**	
	Coeff.	St.Dev	Coeff.	St.Dev	Coeff.	St.Dev	Coeff.	St.Dev
$\Delta P_{p,t-1}$	0.087**	0.037	0.081**	0.037			0.072*	0.037
$\Delta P_{p,t-2}$	0.084**	0.038	0.075*	0.038			0.066*	0.039
$\Delta P_{p,t-1}^{+}$					0.126***	0.045		
$\Delta P_{p,t-1}^{-}$					0.061	0.075		
$\Delta P_{p,t-1}^{+} * D_t$								
$\Delta P_{p,t-1}^{-} * D_t$								
ECT_{t-1}	-0.012**	0.006						
ECT_{t-1}^{+}			0.004	0.011	0.004	0.011	0.013	0.012
ECT_{t-1}^{-}			-0.037**	0.015	-0.042***	0.015	-0.035**	0.015
$ECT_{t-1}^{+} * D_t$							-0.027*	0.016
$ECT_{t-1}^{-} * D_t$							-0.013	0.018
F-Test Short-Run					5.570**			
F-Test Long-Run			6.090**		8.450***		0.13	

Notes: All models control for the constant term α_0 and the $\Delta P_{c,t-i}$ up to 2 lags ($i = 1,2$) as suggested by the SBIC criterion. Results available upon request.
Results of Model (c) AECMSR are robust to the addition of $\Delta P_{p,t-2}^{+}$, $\Delta P_{p,t-2}^{-}$. However, due to the non-significance of coefficients and reduced DoF, we present the more parsimonious estimation. Results of the sensitivity analysis performed with 2 lags are available in Table A8 in the Appendix.
F-Tests on Short- and Long-run asymmetries refer to ECT_{t-1} and $\Delta P_{p,t-1}$ coefficients (positive and negative), respectively

Source: Authors' elaboration

To take into account the structural break previously detected, we estimated an AECM model accounting for a break in the ECT (see panel e and equation (4)). When not accounting for the structural break, the ECT terms' results are consistent with those previously commented. Notably, when the dummy indicating the presence of a structural break is included, the coefficient on positive ECT changes in the post-reform period, i.e., the term $ECT_{t-1}^{+} * D_t$ is statistically significant: when margins are stretching, consumer prices re-adjust to the equilibrium (see the last row of panel d)[16]. The F-test fails to reject the null of symmetry when comparing $ECT_{t-1}^{+} * D_t$ and ECT_{t-1}^{-} (see the last row of panel d).

The estimated models confirm the existence of long-run adjustments, particularly when margin are squeezing, i.e. ECT_{t-1}^{-}. However, despite asymmetric dynamics in price transmission describe the whole period, when accounting for the structural break the price behaviour changes. In the post-reform period, the adjustment towards the equilibrium takes place also when margins are stretching[17].

Following Ben-Kaabia et al. (2005) and Santeramo (2015), we compute the half-lives (Table 3) to conclude on the (average) time (expressed at the same frequency of price series) required to have a ε% price adjustment after an exogenous shock has occurred. Half-lives are computed as follows: $H = \frac{ln(1-\varepsilon)}{ln\,(1+\rho)}$, where ε is the percentage of adjustment, ρ is the strength of adjustment, (i.e., the coefficient α_j), and H represents the months required to observe the ε% adjustments in prices. The half-lives suggest a slight decrease in the speed of adjustment after the break: prices tend to absorb 50% of the shocks in 19 months, but marked differences exist

[16] The model is consistent with different specifications of farmer price, since results do not change when we consider the average EU28 farmer price. Results available upon request.

[17] The ECM Model in panel (a) has been estimated accounting for different demand and supply shifters. Results do not differ from what presented in Table 2, ensuring their robustness (see Table A6 and Table A7 in the Appendix).

when we allow for asymmetric dynamics and account for the policy reform: price adjustments require 20 months when margins squeeze (negative ECT) and up to 26 months when margins stretch.

Table 3: Half-lives for each model specification (values expressed in months)

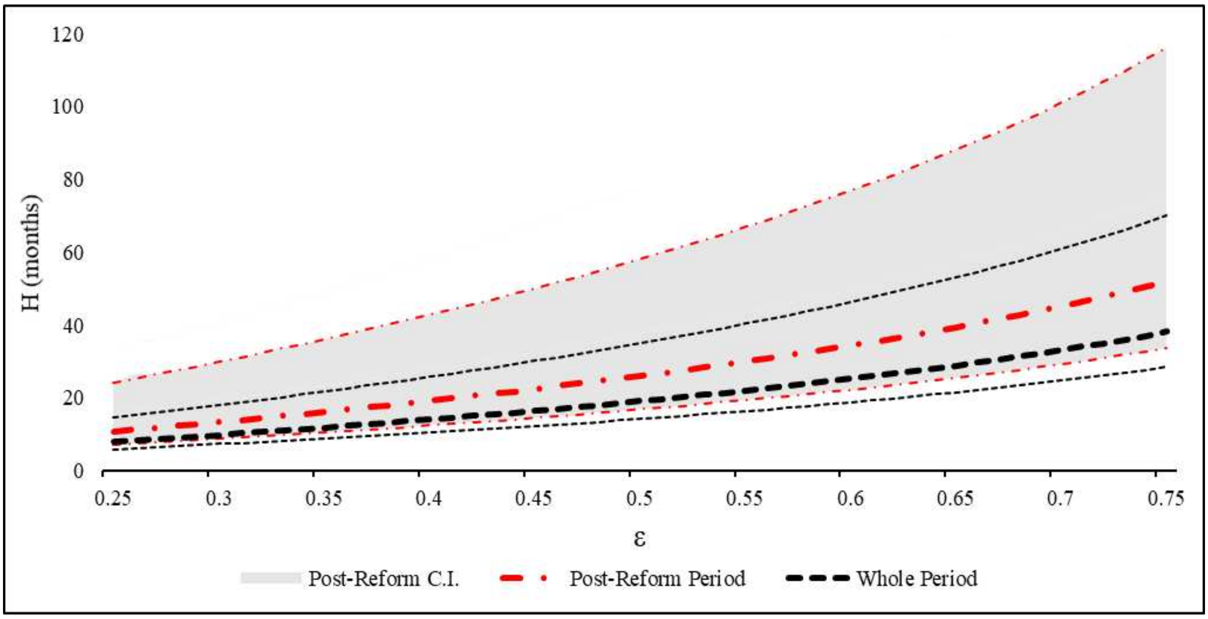
	(a) AECMSR	(b) AECMSB	
ε	ECT_{t-1}^{-}	ECT_{t-1}^{-}	$ECT_{t-1}^{+} * D_t$
0.5	19.1	20.2	26.0
0.7	33.1	34.9	45.2
0.9	63.4	66.9	86.4
0.99	126.8	133.9	172.9

Notes: We calculated half-lives for the statistically significant coefficients in the related model

Source: Authors' elaboration.

Figure 5 illustrates the half-lives related to the AECMSB model (Table 2, panel b), especially regarding the ECT_{t-1}^{-} (whole-period), and the $ECT_{t-1}^{+} * D_t$ (post-Reform period), and their upper and lower bounds. The post-Reform period does not statistically differ from the whole period, as the light-grey shaded area suggests: Whole Period HL falls into the confidence interval of Post-Reform Period.

Figure 5: Half-lives Calculated from the AECMSB with confidence intervals, Whole (ECT_{t-1}^{-}) and Post-Reform ($ECT_{t-1}^{+} * D_t$) Periods



Notes: Confidence intervals were calculated considering the standard errors of each

coefficient estimated in the AECMSB model (see Table 2). Due to very similar results in terms of HL for the ECT_{t-1}^{-} coefficient in AECMSR and AECMSB, only the latter is analyzed.
Source: Authors' elaboration.

Conclusions

The paper explores the price transmission dynamics along the Italian dairy supply chain during a long period in which major reforms occurred, pushing toward a more liberalized market. We investigated price mechanisms and adjustments associated with margin squeezes and stretches. In line with Kinnucan and Forker (1987), who found that government interventions supporting farmers' price favor asymmetric price transmission, we conclude on similar price behavior in the pre-reform period. After the Reform, the fluid milk market seems to pass price signals more efficiently: both positive and negative price changes are transmitted along the supply chain. Our findings are in line with those supported by Cacchiarelli et al. (2016) for the Italian milling industry and by Lee and Gómez (2013) for the international coffee market. An important note has to be devoted to the speed of reaction of prices to exogenous shocks. Prices seem to be more reactive in the pre-reform period than after liberalization. While this is not contrasting with previous studies (*cfr.* Lee and Gómez, 2013), it also signals that fairness (symmetric price adjustments) may come at the cost of market efficiency (slower price transmission). The present analysis reinforces the findings supported by the recent literature on the effects of policy interventions for market liberalization and price stabilization, as envisaged by the risk management measures such as mutual funds and the Income Stabilization Tool (Cordier and Santeramo, 2020), very much promoted in the EU dairy sector (Trestini et al., 2018).

. In addition to the existing evidence, we show that even for a highly perishable product, such as fluid milk, market liberalization favors more symmetric price adjustments along the supply chain. Such results are undoubtedly in line with the envisaged CAP Reform Post-2020 (delayed to 2023), mainly the Farm to Fork strategy. Particular emphasis is devoted to strengthening farmers' position within agricultural markets to retain higher value-added shares and provide fairer supply chains[18]. The Agricultural Markets Task Force (2016), a European Commission's Expert Group, outlined how CAP became far more market-oriented, integrating EU agricultural markets into global value chains, exposing fragmented agricultural producers as the main shock-absorbers. Therefore, for a better assessment of farmers' position within the supply chain, the Group calls for more transparent and coordinated data collection, especially concerning prices at different steps of the agricultural chains. Future research should be devoted to investigating the vertical price transmission along the global value chains, particularly in view of the influence exerted by countries' specialization in production, trade relationships, and the global challenges (e.g., climate change) (Santeramo et al., 2021).

We recognize potential limitations of the present study: the price data, national average price indexes do not allow us to account for quality characteristics; second, the monthly frequency may hinder shorter-term price dynamics; third, while it is common to assume no substitution effects on the supply side, a liter of milk - subjected to technical restrictions based on milk and fat content – can be processed into different dairy products and the profit-maximizing processors consider the prices of different dairy products into account when deciding to process the raw product[19].

[18] For more details see the Communication from the Commission "A Farm to Fork Strategy for a fair, healthy and environmentally-friendly food system", available at https://eur-lex.europa.eu/resource.html?uri=cellar:ea0f9f73-9ab2-11ea-9d2d-01aa75ed71a1.0001.02/DOC_1&format=PDF

[19] We are grateful to an anonymous referee for such suggestion.

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Appendix

Table A1: Unit-Root Tests for Pc and Pp both in levels and first difference.

Unit-Root Test	Lags	Stat	5% C.V.	IC	Result	Lags	Stat.	5% C.V.	IC	Result
			Pc						**ΔPc**	
DF-GLS	4	-1.859	-2.911	Ng-Perron	I(1)	0	-6.072	-2.937	Ng-Perron	I(0)
(w/Trend)*	1	-1.125	-2.936	SBIC, MAIC		3	-3.816	-2.936	SBIC, MAIC	
PP*	4	-2.411	-2.883	Newey-West	I(1)	4	-5.696	-2.883	Newey-West	I(0)
PP (w/Trend)*	4	-1.178	-3.437	Newey-West		4	-5.933	-3.437	Newey-West	
KPSS (w/Trend)	4	0.47	0.148		I(1)	8	0.051	0.148		I(0)
KPSS	4	3.94	0.462			8	0.319	0.462		
			Pp						**ΔPp**	
DF-GLS	11	-2.116	-2.837	Ng-Perron	I(1)	10	-3.929	-2.849	Ng-Perron	I(0)
(w/Trend)*	2	-2.162	-2.928	MAIC		8	-3.047	-2.937	MAIC	
PP (w/Trend)*	4	-1.691	-2.883	Newey-West	I(1)	4	-7.957	-3.437	Newey-West	I(0)
PP*	4	-1.248	-3.437	Newey-West		4	-7.848	-2.883	Newey-West	
KPSS (w/ Trend)	4	0.264	0.148		I(1)	4	0.101	0.148		I(0)
KPSS	4	3.727	0.462			4	0.224	0.462		

*Maximum lag-length selection set to 12

Source: Authors' elaboration

The Johansen et al. (2000) cointegration test we applied is as follows:

$$\Delta P_t = \alpha\big(\beta' P_{t-1} + \delta_1(t-1)D_{1.t} + \delta_2(t-1)D_{2,t}\big) + \sum_{i=1}^{k-1}\Gamma_i \Delta P_{t-i} + \sum_{i=1}^{k}\theta_i D_{1,t-i} + \varepsilon_t$$

Where ΔP_t is the vector of our price series; $D_{1.t}$ is a dummy variables which takes the value 1 whenever $t>T_1$ and zero otherwise; $D_{2,t} = 1 - D_{1,t}$; Γ_i and θ_i are matrices of short-run parameters; δ_1, δ_2 are parameter vectors referred to the intercepts of the two regimes.

Marginal Model for Testing Exogeneity for the Pp:

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1}\mathit{\Gamma}_i\,\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\vartheta_i\,\Delta P_{p,t-i} + v_t \qquad \text{(a)}$$

the $\hat{v}_{t-1}$ were estimated and plugged into the ECR

$$\Delta P_c = \alpha_0 + \alpha_1\big(P_{c,t-1} - \beta_0 - \beta_1 P_{p,t-1}\big) + \sum_{i=1}^{k-1}\mathit{\Gamma}_i\,\Delta P_{c,t-i} + \gamma_1 \Delta P_p + \sum_{i=1}^{k-1}\vartheta_i\,\Delta P_{p,t-i} + \delta_1\hat{v}_{t-1} + \mu_t \qquad \text{(b)}$$

as well as into the AECR

$$\Delta P_c = \alpha_0 + a_1^{+}ECT_{t-1}^{+} + a_2^{-}ECT_{t-1}^{-} + \sum_{i=1}^{k-1}\mathit{\Gamma}_i\,\Delta P_{c,t-i} + \gamma_1 \Delta P_p + \sum_{i=1}^{k-1}\vartheta_i\,\Delta P_{p,t-i} + \delta_1\hat{v}_{t-1} + \varepsilon_t \qquad \text{(c)}$$

The residuals from the marginal model result in being non-significant in both cases, and, hence, prevents the rejection of the null hypothesis of weak-exogeneity to short-run parameters. Aimed at testing also the weak-exogeneity of Pp regarding the long-run parameters, we tested for the statistical significance of the $ECT_{t-1}, ECT_{t-1}^{+}, ECT_{t-1}^{-}$ into the marginal model, such as:

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1}\mathit{\Gamma}_i\,\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\psi_i\,\Delta P_{p,t-i} + \alpha_1\big(P_{c,t-1} - \beta_0 - \beta_1 P_{p,t-1}\big) + v_t \qquad \text{(d)}$$

and:

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1}\mathit{\Gamma}_i\,\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\psi_i\,\Delta P_{p,t-i} + a_1^{+}ECT_{t-1}^{+} + a_2^{-}ECT_{t-1}^{-} + v_t \qquad \text{(e)}$$

All the ECT added in the two models resulted being non-significant. Thus, we cannot reject the null hypothesis of Pp weak-exogeneity from long-run parameters (Table A5).

Table A2: Marginal Model Estimates for Short (a,b, and c) and Long (d, and e) Parameters.

	Coef.	**Std. Err.**	**t**	**P>t**
	Marginal Model Estimates (a)			
Γ_ι*	0.1503525	0.1123403	1.34	0.182
ϑ_ι*	0.4932108	0.0686206	7.19	0.000
α_0	0.0003261	0.0003793	0.86	0.391
*SIC indicates one lag should be included				
	Estimates from the Marginal Model with the ECT (d)			
ϑ_ι	0.4980574	0.0700114	7.11	0
$\boldsymbol{\alpha_1}$	**0.0046224**	**0.0125035**	**0.37**	**0.712**
Γ_ι	0.1626874	0.1174296	1.39	0.168
α_0	0.000298	0.0003877	0.77	0.443
	Estimates from the AEC Marginal Model (e)			
ϑ_ι	0.4870199	0.0706924	6.89	0
$\boldsymbol{\alpha_1^+}$	**0.0257962**	**0.0229768**	**1.12**	**0.263**
$\boldsymbol{\alpha_1^-}$	**-0.0247918**	**0.0295569**	**-0.84**	**0.403**
Γ_ι	0.1391981	0.1193005	1.17	0.245
α_0	-0.0002894	0.0006605	-0.44	0.662

	Coef.	**Std. Err.**	**t**	**P>t**
	Weak-Exogeneity Test in the ECR (b)			
ϑ_ι	0.156332	0.071532	2.19	0.03
γ_1	0.137902	0.033876	4.07	0
α_1	-0.01422	0.006003	-2.37	0.019
Γ_ι	0.49567	0.058995	8.4	0
$\boldsymbol{\delta_\iota}$	**-0.13265**	**0.070574**	**-1.88**	**0.062**
α_0	0.000566	0.000183	3.09	0.002
	Weak-Exogeneity Test in the AECR (c)			
Γ_ι	0.484048	0.059137	8.19	0
α_1^+	0.000526	0.01069	0.05	0.961
α_1^-	-0.03566	0.014202	-2.51	0.013
ϑ_ι	0.137234	0.072122	1.9	0.059
γ_1	0.134943	0.033767	4	0
$\boldsymbol{\delta_\iota}$	**-0.11782**	**0.070812**	**-1.66**	**0.098**
α_0	0.000155	0.000307	0.51	0.613

Source: Authors' elaboration

Table A3: Zivot and Andrews Test for Structural Breaks

		Pc		
Lags	Break	t-stat	10%	Results
1	September 2007	-3.819	-4.58	Accept the null of I(1) with a structural break
		Pp		
Lags	Break	t-stat	10%	Results
2	June 2007	-4.167	-4.58	Accept the null of I(1) with a structural break

Source: Authors' elaboration

Table A4: Further Cointegration Tests restricting a trend and a structural break to the cointegratin space

Rank	**Trace**	**P-Value**
	Restricted trend	
0	13.220	0.723
1	5.020	0.601
	Restricted trend and dummy variable (i.e., structural break)	
0	22.280	0.226
1	5.540	0.643

Note: Number of lags included: 2 (based on the Schwartz Information Criteria)

Source: Authors' elaboration

Table A5: Granger Causality Tests on VAR(1) and VAR(2)

		VAR(2)		
Equation	Excluded	χ^2	df	p-value
ΔPc	**ΔPp**	18.285	2	0.000
ΔPp	**ΔPc**	2.108	2	0.348
		VAR(1)		
ΔPc	**ΔPp**	15.335	1	0.000
ΔPp	**ΔPc**	1.8102	1	0.178

Source: Authors' elaboration

Table A6: Demand and Supply Shifters

Label	**Short Description**	**Source**	**Frequency**	**Unit-Root***
Z^1_r	Unemployment Rate - Seasonally Adjusted	Eurostat	Monthly	Yes
Z^2_r	GDP and main components (output, expenditure and income)	Eurostat	Quarterly	Yes
Z^1_p	Industrial Inputs Price	IMF	Monthly	Yes
Z^2_p	Energy Price Index	IMF	Monthly	Yes
P_{EU}	Weighted Average EU Milk Price	MMO (EC)	Monthly	Yes

*We tested for the rpesence of unit root by ADF-GLS, PP, and KPSS tests. Results are available upon request

Source: Authors' elaboration

Table A7: ECM Model Results for Demand and Supply Shifters

ΔCp	Coef.	Std. Err.	t	P>t	ΔCp	Coef.	Std. Err.	t	P>t
ECM and P_{EU}					ECM and Z^1_r (Unemployment Rate)				
ΔCp_{t-1}	0.558	0.057	9.76	0.000	ΔCp_{t-1}	0.564	0.057	9.800	0.000
ECT_{t-1}	-0.015	0.006	-2.61	0.010	ECT_{t-1}	-0.015	0.006	-2.580	0.011
ΔP_{EUt-1}	0.010	0.006	1.6	0.111	$\Delta Z^1_{r,t}$	-0.006	0.006	-0.920	0.360
ΔPp_{t-1}	0.088	0.037	2.34	0.020	ΔPp_{t-1}	0.113	0.034	3.320	0.001
a_0	0.000	0.000	3.65	0.000	a_0	0.000	0.000	3.440	0.001
ECM and Z^2_r (GDP)					ECM and Z^2_p (Energy Price Index)				
ΔCp_{t-1}	0.557	0.057	9.750	0	ΔCp_{t-1}	0.559	0.057	9.770	0.000
ECT_{t-1}	-0.015	0.006	-2.550	0.012	ECT_{t-1}	-0.015	0.006	-2.610	0.010
$\Delta Z^2_{r,t}$	0.051	0.031	1.680	0.094	$\Delta Z^2_{p,t}$	0.003	0.002	1.420	0.158
ΔPp_{t-1}	0.113	0.034	3.330	0.001	ΔPp_{t-1}	0.109	0.034	3.170	0.002
a_0	0.000	0.000	2.960	0.003	a_0	0.000	0.000	3.450	0.001
ECM and Z^1_p (Input Cost)									
ΔCp_{t-1}	0.560	0.057	9.730	0.000					
ECT_{t-1}	-0.015	0.006	-2.510	0.013					
$\Delta Z^1_{p,t}$	-0.000	0.004	-0.020	0.985					
ΔPp_{t-1}	0.115	0.034	3.360	0.001					
a_0	0.000	0.000	3.420	0.001					

Note: Estimating together supply and demand shifters in the same model brought to non-significant estimates

Source: Authors' elaboration

Table A8: Model results for the ASRM with 1 and 2 lags, and the AECMSR with two lags and asymmetries

	(a.1) ASRM		**(a.2) ASRM**		**(a.3) AECMSR**	
	Coeff.	St.Dev	Coeff.	St.Dev	Coeff.	St.Dev
$\Delta P_{p,t-1}$						
$\Delta P_{p,t-2}$						
$\Delta P^{+}_{p,t-1}$	0.149***	0.045	0.110**	0.052	0.098*	0.052
$\Delta P^{-}_{p,t-1}$	0.096	0.076	0.076	0.077	0.051	0.077
$\Delta P^{+}_{p,t-2}$			0.079	0.053	0.062	0.053
$\Delta P^{-}_{p,t-2}$			0.124	0.077	0.093	0.077
$\Delta P^{+}_{p,t-1} * D_t$						
$\Delta P^{-}_{p,t-1} * D_t$						
ECT_{t-1}						
ECT^{+}_{t-1}					0.004	0.011
ECT^{-}_{t-1}					-0.036***	0.015
$ECT^{+}_{t-1} * D_t$						
$ECT^{-}_{t-1} * D_t$						
F-Test Short-Run (lag 1; lag 2)					3.380*; 3.390*	
F-Test Long-Run					6.050***	

Source: Authors' elaboration