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

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During the last 2 decades, two policy reforms—the Fischler Reform and the Common Market Organization Reform—have pushed the EU dairy sector toward economic liberalization. These changes affected the EU supply chains at different levels, altering the mechanisms of vertical price transmission. Against this backdrop, 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.

Key words: asymmetries, CAP reform, dairy sector, error correction model, Fischler Reform, structural break

Introduction

The Italian dairy industry, especially the fluid milk sector, has experienced relevant changes in the last decades. Searching for liberalization is not new in the EU (see, e.g., the McSharry Reform in 1992) (Swinbank, 1993; Coleman and Tangermann, 1999); the 2003 reform enhanced the farm-gate competitiveness and developed a more market-driven dairy industry, mainly due to the abolishment of the quota regime (Henke et al., 2018). It has also stiffened 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 (see 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 speed of price transmission proxy the degree of integration of the supply chain and inform on market efficiency, a key feature for planning strategic and structural market interventions (Goodwin and Holt, 1999; Serra and Goodwin, 2003).

The literature on price transmission is vast, but the majority of studies have focused on spatial/horizontal price transmission (Awokuse, 2007; Cioffi, Santeramo, and Vitale, 2011; Santeramo and Cioffi, 2012; Chen and Saghaian, 2016; Durborow et al., 2020; Hatzenbuehler, Du, and Painter, 2021; Santeramo, 2022), with a few exploring the dairy sector (Fabiosa et al., 2005; Kempen et al., 2011; Bergmann, O'Connor, and Thümmel, 2015; Hillen and von Cramon-Taubadel, 2019). Applications of the VPT analysis to the food supply chains are also numerous

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(e.g., Ben-Kaabia and Gil, 2007; Boetel and Liu, 2010; Li and Sexton, 2013; Ahn and Lee, 2015; Antonioli et al., 2019), but only a few have analyzed the impacts of the CAP reforms (e.g., Cacchiarelli, Lass, and Sorrentino, 2016; Rezitis and Pachis, 2018), especially on the milk supply chain.[1]

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 that may explain observed changes in bargaining power and price transmission dynamics (Rama, 2019). The unbalanced market power, which favors retailers; the rigidity of the supply, as implied by policy measures; and the low storability of the fresh fluid milk are likely to have favored asymmetric price transmission dynamics.[2] The CAP reform has liberalized the market and likely favored a whole, symmetric, fast transmission of market signals.

Using processors and consumers' prices for fresh fluid milk from January 2000 to August 2016, we explore how the 2003 Common Agricultural Policy (CAP) reform, intended to liberalize EU dairy markets, has influenced VPT dynamics in the context of Italy. We assess the nature of price transmission (i.e., cost-push, demand-pull, or feedback system), the degree the symmetry, and the speed at which shocks are passed through.

The Dairy Sector: Policy and Economic Facts

The CAP reform of 2003—referred to as Fischler Reform—liberalized the EU dairy sector "by reducing price support and creating direct income support" (European Court of Auditors, 2009, p. 13). After the reform, the milk target price were removed, intervention prices for dairy products were lowered, and national milk quotas were dismantled (Gohin and Latruffe, 2006). In 2007, the new Single Common Market Organization also had a significant impact on the European milk market, setting export subsidies for milk to 0 (with exceptions made for exports made from January to November 2009) and creating new intervention prices for butter and skimmed milk powder (Meijerink and Achterbosch, 2013). On March 1, 2015, the quota system,[3] introduced in 1984 to regulate the supply surplus and sustain farm-level prices in the EU,[4] was dismantled entirely (Giles, 2015).[5]

The total support estimate (TSE), an Organisatio for Economic Cooperation and Development (2021) indicator that combines all agriculture-related public expenses, and the producer support estimate (PSE), which proxies the direct support paid to agricultural producers, have declined tremendously since 2007 (Figure 1). These dynamics suggest that the CAP has moved to a far less intervening policy framework: Since 2007, the EU dairy sector has received almost no support.

The Italian dairy industry is valued at around €15 billion (11.4% of Italy's food industry), embracing more than 2,000 dairy firms and 30,000 employees and accounting for more than 7% of total EU-28 production, although remaining a net importer of dairy. Raw milk is sold as

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

[2] According to Rama (2019), the CR4 for dairy processors increased from approximately 12% to 15% during the 2000–2016 period, while food retailers experienced an increase in CR4 from 38% to 45% (Autorità Garante della Concorrenza e del Mercato, 2013; Mediobanca, 2020), maintaining market power. Farm populations 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 head per farm has doubled since 2000, to approximately 45 head per farm in 2016).

[3] For more details on the aforementioned policy interventions, 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 is the amount of rent generated from a restriction on supply (Tonini and Dominguez, 2009).

[5] The abolition of the quota was preceded, in 2008, by phasing-out measures: Quotas increased by 1%–2% for 5 consecutive years (exceptions were made for Italy, for which there was a 5% increase in the 2009–2010 campaign) (European Commission, 2010).

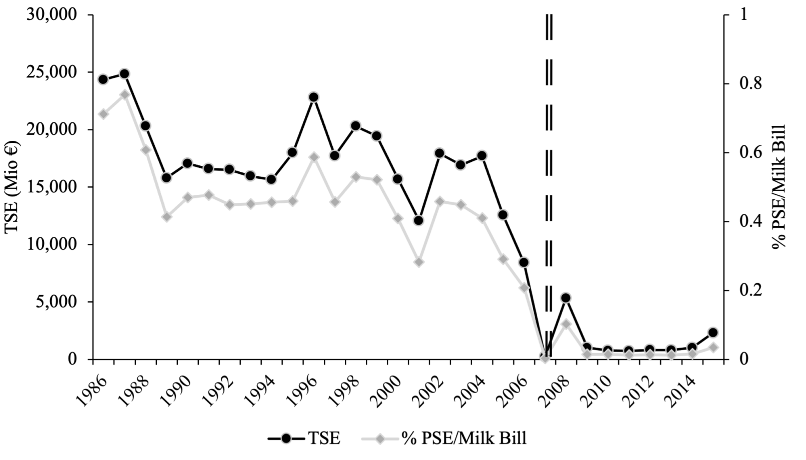

Figure 1. Total Support Estimate and Producer Support Estimate in the EU Dairy Sector

Source: Authors' elaboration on OECD data.

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 (Rama, 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 (Rama, 2019).

Literature Review

Gardner's (1975) seminal paper on price transmission warns of 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), but they tend to be more common than expected (Peltzman, 2000; Bakucs, Fałkowski, and Fertő, 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) find evidence of asymmetries for nonperishable 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 (2015) find that highly perishable fruits are characterized by negative asymmetries, while the opposite is true for less perishable ones. Santeramo (2015) supports Kim and Ward's (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 price transmission 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, Santeramo, and Vitale (2011) conclude that the system of quotas and tariffs in the fruit and vegetable sector alters 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, Lass, and Sorrentino (2016) investigate the impacts of the midterm CAP reform and concluded on the symmetry (asymmetry) of farm–wholesaler (wholesaler–retailer) prices. Finally, McCorriston (2015) concludes that consumers' reaction to positive and negative price movements may induce asymmetric VPT, a conclusion supported by Biden, Ker, and Duff (2020) with evidence for the dairy sector.

In summary, price transmission is informative about 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 are accounted for using an error correction model (ECM) (Engle and Granger, 1987):

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

where $P_{c,t-1} - \alpha_0 - \beta_1 P_{p,t-1} = ECT_{t-1}$ represents the error correction term (ECT). Equation (1) assumes producer price to be exogenous and the cost-push mechanism to lead the PT dynamics (see Ben-Kaabia and Gil, 2007; Santeramo and von Cramon-Taubadel, 2016). The assumption is empirically tested (see footnote 14).

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 ECT (Granger and Lee, 1989):

$$\Delta P_{c,t} = \alpha_0 + \alpha_1 ECT^{+}_{t-1} + \alpha_2 ECT^{-}_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \Delta P_{p,t-i} + \varepsilon_t, \tag{2}$$

where $ECT^{+}_{t-1} = \begin{cases} ECT_{t-1} \text{ if } ECT_{t-1} > 0 \\ 0 \text{ otherwise} \end{cases}$ and $ECT^{-}_{t-1} = \begin{cases} ECT_{t-1} \text{ if } ECT_{t-1} < 0 \\ 0 \text{ otherwise} \end{cases}$.

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:

$$\begin{aligned} \Delta P_{c,t} = {} & \alpha_0 + \alpha_1 ECT^{+}_{t-1} + \alpha_2 ECT^{-}_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_{1i} \Delta P^{+}_{p,t-i} \\ & + \sum_{i=1}^{k-1} \psi_{2i} \Delta P^{-}_{p,t-i} + \varepsilon_t. \end{aligned} \tag{3}$$

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) test and modeled via an AECM with a structural break in the ECT:

$$\begin{aligned} \Delta P_{c,t} = {} & \alpha_3 + (1 - D_t) \cdot \left(\alpha_1 ECT^{+}_{t-1} + \alpha_2 ECT^{-}_{t-1}\right) + D_t \cdot \left(\alpha_3 ECT^{+}_{t-1} + \alpha_4 ECT^{-}_{t-1}\right) \\ & + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \Delta P_{p,t-i} + \varepsilon_t, \end{aligned} \tag{4}$$

where $D_t = \begin{cases} 1 \text{ if } t \geq t_{bp} \\ 0 \text{ if } t < t_{bp} \end{cases}$ 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.

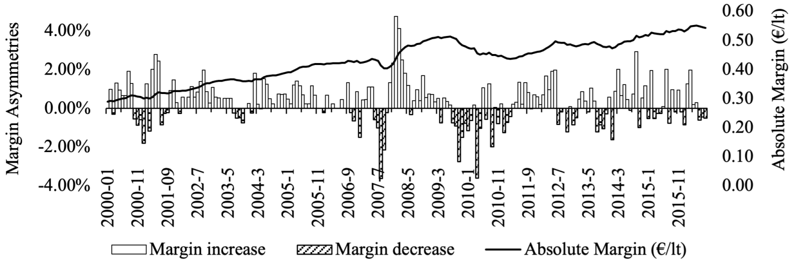

Figure 2. Italian Milk Margin Trend

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

Preliminary Analysis

In the post-reform period, prices became highly volatile:[6] Consumer prices increased by 28% (volatility, σ_t, was 0.33 in the pre-reform period, $\sigma_{Jan\ 00-Jul\ 07}$, and 0.42 in the post-reform period, $\sigma_{Aug\ 07-Aug\ 16}$).[7] The processor price's volatility rose by 83% ($\sigma_{\text{Jan }00-Jul\ 07} = 0.37$, $\sigma_{\text{Aug }07-\text{Aug }16} = 0.69$), suggesting that retailers do not entirely pass the changes in industrial processor price to consumers.

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). Figure 2 shows that positive margin changes were more frequent but lower in magnitude (€0.37/l on average) in the prebreak period 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. Further, more than 60% of the negative changes in margins occurred in the post-2007 period, characterized by a more symmetric evolution of margin changes.

In the pre-reform period, positive margins are 2.8 times the frequency of decreasing margins; the post-reform ratio is 1.4 and the periods of margin decrease doubled, from 23 to 44. Consumer price (P_c) upward movements shrank by 17% post-reform, while P_c reductions are 15 times more likely than in the pre-reform scenario. Processor prices (P_p) 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%) relative to the pre-reform period and more evenly distributed, signaling a more balanced bargaining power along the chain.

Data and Results

We use monthly data from January 2000 to August 2016 on the price indices (August 2010 = 100 of the fluid milk supply chain (Figure 4). The price paid by retailers to fluid milk processors (P_p) and the price applied by the retailers to consumers (P_c) were collected from the Italian National Institute of Statistics (ISTAT).[8]

The unit-root tests with constant suggest that the series are integrated of order 1 ($I(1)$).[9] The Zivot–Andrews test points to structural breaks in June 2007 (processor price) and September 2007

[6] 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.

[7] Both the pre- and post-reform periods refer to the structural break empirically detected in August 2007.

[8] Notably, given that industrial processing (i.e., pasteurization and packaging) does not significantly transform the fluid milk collected from farmers, our processor price is a valid proxy of the farm-gate price.

[9] We employed diverse unit-root tests, all of which clearly point to $I(1)$ series (results available in Table S1 in the online supplement, www.jareonline.org).

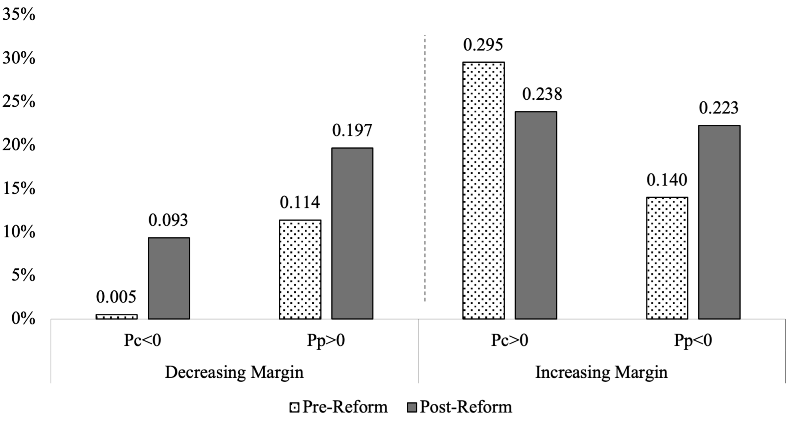

Figure 3. Margin Changes and Price Dynamics in Pre- and Post-Reform Periods

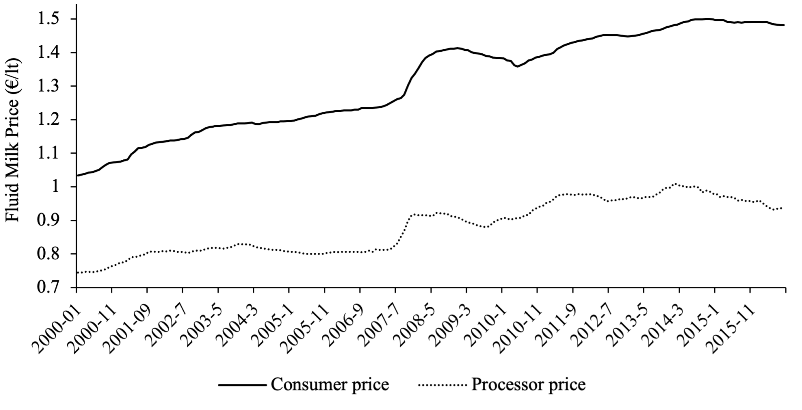

Figure 4. Consumer and Processor Prices of Fresh Fluid Milk in the Italian Market, 2000–2016

Notes: For a more comprehensive graphical representation, we multiplied the two indices by the absolute nominal price of fluid milk observed in August 2010, which is the base on which the two indices rely. The absolute market price for both supply chain levels 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).

(consumer price):[10] 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,[11] we obtain $P_c = 0.206P_p + 0.149D_t + 3.576$, where D_t is a dummy variable taking the value of 1 after August 2007, when the structural break occurs, and 0 otherwise.[12] The weak-exogeneity test allows us to draw conclusions about the exogeneity of the processor price, signaling a cost-push mechanism, in line with most studies on PT in agricultural markets.[13]

[10] See Table S3 in the online supplement for detailed results.

[11] 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 from the authors upon request.

[12] We also investigate the presence of cointegration restricting a trend into the cointegrating space, both with and without a structural break, and conclude that there is no cointegrating relationship. See Table S4 in the online supplement for further details.

[13] In order to reinforce the exogeneity assumption, we ran a Granger causality test, leading to the exogeneity of processor price. The marginal model (see von Cramon-Taubadel, 1998) estimated for the sake of model consistency points to exogenous processor price. Accordingly, we set P_c as the independent variable. See Tables S2 and S5 in the online supplement for more details.

Table 1. Results of Cointegration Test with Structural Breaks

Rank	Trace	p-Value
0	27.832	0.033
1	8.224	0.266

Notes: The constant is restricted to the cointegrated space, together with the dummy variable D_t taking a value of 1 when $t \geq 2007:08$ (i.e., the structural break), and 0 otherwise. Two lags are included (Schwartz information criterion).

The preliminary results on the ECM model (see equation 1) suggest that a long-run relationship links prices (see column 1 of Table 2),[14] where about 12% of a change in processor prices is transmitted to consumer prices. This suggests a sluggish adjustment of the system toward the equilibrium once a shock occurs.

The asymmetric short-run model (ASRM), specified as

$$\Delta P_{c,t} = \alpha_0 + \sum_{i=1}^{k-1} \Gamma_i \Delta Pc_{t-i} + \sum_{i=1}^{k-1} \psi_{1i} \Delta P^{+}_{p,t-i} + \sum_{i=1}^{k-1} \psi_{2i} \Delta P^{-}_{p,t-i} + \varepsilon_t$$

(Table S8, column 1), suggests that positive changes in processor prices influence consumer prices: When processor prices increase (and margins 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, column 2): When margins squeeze (negative ECT), consumer prices react toward long-run equilibrium. The F-test on asymmetry suggests that price adjustments are asymmetric. Estimates in column 3 include asymmetric dynamics in both the ECT and the short run (see equation 3). Results indicate that processor price increases and negative ECT (i.e., margin squeeze) influence consumer prices, consistent with results from column 2. Overall, we found asymmetries in both long- and short-run price adjustments, as the F-test suggests in the last two rows of column 3, a result 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).

To account for the structural break previously detected, we estimated an AECM model accounting for a break in the ECT (see column 4 of Table 2 and equation 4). When not accounting for the structural break, the results of the ECT terms are consistent with those previously noted. In particular, 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 $ECT^{+}_{t-1} \times D_t$) is statistically significant: When margins are stretching, consumer prices readjust to the equilibrium (see the last row of column 4 of Table 2).[15] The F-test fails to reject the null of symmetry when comparing $ECT^{+}_{t-1} \times D_t$ and ECT^{-}_{t-1} (see the last row of column 4). 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 that describe the whole period, price behavior changes when accounting for the structural break. In the post-reform period, the adjustment toward the equilibrium also occurs when margins are stretching.[16]

Following Santeramo (2015), we compute the half-lives (HL) (Table 3) to determine the (average) time (expressed at the same frequency of price series) required to have a $\varepsilon\%$ price adjustment after an exogenous shock has occurred.

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

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

[16] The ECM model in column 1 of Table 1 has been estimated accounting for different demand and supply shifters. Results do not differ from what presented in Table 2, ensuring their robustness (see Tables S6 and S7 in the online supplement).

Table 2. Results from the Estimated Models

	ECM	AECM	AECMSR	AECMSB
	1	2	3	4
$\Delta P_{p,t-1}$	0.087**	0.081**		0.072*
	(0.037)	(0.037)		(0.037)
$\Delta P_{p,t-2}$	0.084**	0.075*		0.066*
	(0.038)	(0.037)		(0.039)
$\Delta P^{+}_{p,t-1}$			0.126***	
			(0.045)	
$\Delta P^{-}_{p,t-1}$			0.061	
			(0.075)	
ECT_{t-1}	−0.012**			
	(0.006)			
ECT^{+}_{t-1}		0.004	0.004	0.013
		(0.011)	(0.011)	(0.012)
ECT^{-}_{t-1}		−0.037**	−0.042***	−0.035**
		(0.015)	(0.015)	(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 nonsignificance of coefficients and reduced DoF, we present the more parsimonious estimation. Results of the sensitivity analysis performed with 2 lags are available in Table S8 in the online supplement. F-tests on short- and long-run asymmetries refer to ECT_{t-1} and $\Delta P_{p,t-1}$ coefficients (positive and negative), respectively.

Table 3. Half-Lives for Each Model Specification (in months)

	AECMSR	AECMSB	
$\boldsymbol{\varepsilon}$	$\boldsymbol{ECT_{t-1}}^{-}$	$\boldsymbol{ECT_{t-1}}^{-}$	$\boldsymbol{ECT_{t-1}}^{+} \times \boldsymbol{D_t}$
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 calculate half-lives for the statistically significant coefficients in the related model.

Half-lives are computed as follows: $H = \frac{\ln(1-\varepsilon)}{\ln(1+\rho)}$, where ε is the percentage of adjustment, *rho* 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 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.

Figure 5 illustrates the half-lives related to the AECMSB model (Table 2, column 4), especially regarding the ECT^{-}_{t-1} (whole period) and the $ECT^{+}_{t-1} \times D_t$ (post-reform period) and their upper and

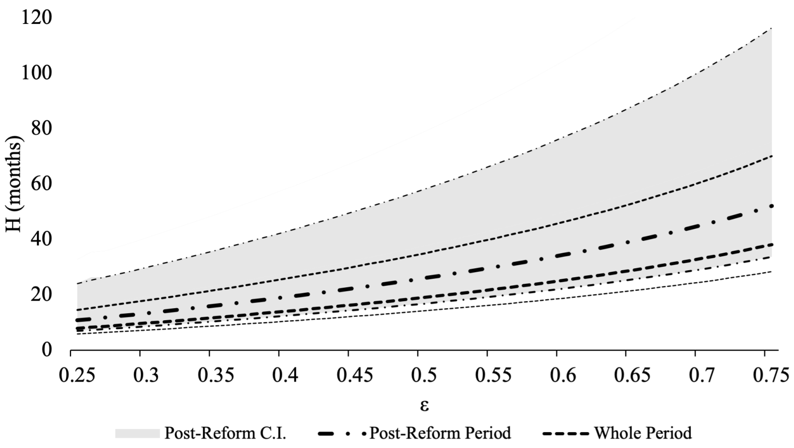

Figure 5. Half-Lives Calculated from the AECMSB with Confidence Intervals, Whole (ECT^{-}_{t-1}) and Post-Reform ($ECT^{+}_{t-1} \times D_t$) Periods

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.

lower bounds. The post-reform period does not statistically differ from the whole period, as the area shaded light grey suggests: Whole-period HL falls into the confidence interval of the post-reform period.

Conclusions

This paper explored the price transmission dynamics along the Italian dairy supply chain over a long period during which major reforms occurred, pushing toward a more liberalized market. We investigated price mechanisms and adjustments associated with margin squeezes and stretches. Like Kinnucan and Forker (1987), who found that government interventions supporting farmers' price favor asymmetric price transmission, we find 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 support those by Cacchiarelli, Lass, and Sorrentino (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 does not contradict previous studies (see 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 market liberalization favors more symmetric price adjustments along the supply chain, even for a highly perishable product such as fluid milk. Such results are undoubtedly in line with the envisaged CAP reform post-2020 (delayed to 2023), particularly 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 (for more details, see European Commission, 2020). The Agricultural Markets Task Force (2016), a European Commission 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 investigate vertical price transmission along the global value chains, particularly in view of the influence exerted by countries' specialization in production, trade relationships, and global challenges (e.g., climate change) (Santeramo, Miljkovic, and Lamonaca, 2021).

We recognize potential limitations of the present study. National average price indices do not allow us to account for quality characteristics, and monthly frequency may hinder examination of shorter-term price dynamics. Additionally, 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 profit-maximizing processors consider the prices of different dairy products when deciding to process the raw product.[17]

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Online Supplement: On Policy Interventions and Vertical Price Transmission: The Italian Milk Supply Chain Case

Federico Antonioli and Fabio Gaetano Santeramo

The GLS-ADF test (Elliott, Rothenberg, and Stock, 1996), the KPSS test (Kwiatkowski et al., 1992), and the PP test (Phillips and Perron, 1988), are applied.

Table S1. Unit-Root Tests for P_c and P_p Both in Levels and First Difference

Unit-Root Tests	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)[a]	1	-1.125	-2.936	SBIC, MAIC		3	-3.816	-2.936	SBIC, MAIC	
PP[a]	4	-2.411	-2.883	Newey-West	$I(1)$	4	-5.696	-2.883	Newey-West	$I(0)$
PP (w/Trend)[a]	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)[a]	2	-2.162	-2.928	MAIC		8	-3.047	-2.937	MAIC	
PP (w/Trend)[a]	4	-1.691	-2.883	Newey-West	$I(1)$	4	-7.957	-3.437	Newey-West	$I(0)$
PP[a]	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		

Notes: [a] Maximum lag-length selection set to 12
Source: Authors' elaboration

The Johansen, Mosconi, and Nielsen (2000) cointegration test we applied is as follows:

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

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 0 otherwise; $D_{2,t} = 1 - D_{1,t}; \Gamma_i$ 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

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1}\Gamma_i\Delta P_{c,t-i} + \sum_{i=1}^{k-1}\vartheta_i\Delta P_{p,t-i} + v_t \tag{S1}$$

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

$$\begin{aligned}\Delta P_c = {} & \alpha_0 + \alpha_1\left(P_{c,t-1} - \beta_0 - \beta_1 P_{p,t-1}\right) + \sum_{i=1}^{k-1}\Gamma_i\Delta P_{c,t-i} \\ & + \gamma_1\Delta P_p + \sum_{i=1}^{k-1}\vartheta_i\Delta P_{p,t-i} + \delta_1\hat{v}_{t-1} + \mu_t\end{aligned} \tag{S2}$$

as well as into the AECRjoh

$$\Delta P_c = \alpha_0 + a_1^+ ECT_{t-1}^+ + a_2^- ECT_{t-1}^- + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \gamma_1 \Delta P_p + \sum_{i=1}^{k-1} \vartheta_i \Delta P_{p,t-i} + \delta_1 \hat{v}_{t-1} + \varepsilon_t \tag{S3}$$

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:

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \Delta P_{p,t-i} + \alpha_1 \left(P_{c,t-1} - \beta_0 - \beta_1 P_{p,t-1}\right) + v_t \tag{S4}$$

and:

$$\Delta P_p = \alpha_0 + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{c,t-i} + \sum_{i=1}^{k-1} \psi_i \Delta P_{p,t-i} + a_1^+ ECT_{t-1}^+ + a_2^- ECT_{t-1}^- + v_t \tag{S5}$$

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 S5).

Table S2. Marginal Model: Estimates for Short (Panels A, B, and C) and Long (Panels D and E) Parameters

	Coef.	Std. Err.	t	P>t
Panel A. Marginal Model Estimates				
Γ_l^*	0.1503525	0.1123403	1.34	0.182
ϑ_1^*	0.4932108	0.0686206	7.19	0.000
α_0	0.0003261	0.0003793	0.86	0.391
Panel B. Weak-Exogeneity Test in the ECR				
ϑ_1	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_1	0.49567	0.058995	8.4	0
$\boldsymbol{\delta}_1$	-0.13265	0.070574	-1.88	0.062
α_0	0.000566	0.000183	3.09	0.002
Panel C. Weak-Exogeneity Test in the AECR				
Γ_l	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_1	0.137234	0.072122	1.9	0.059
γ_1	0.134943	0.033767	4	0
$\boldsymbol{\delta}_1$	-0.11782	0.070812	-1.66	0.098
α_0	0.000155	0.000307	0.51	0.613
Panel D. Estimates from the Marginal Model with the ECT				
ϑ_1	0.4980574	0.0700114	7.11	0
$\boldsymbol{\alpha}_1$	0.0046224	0.0125035	0.37	0.712
Γ_1	0.1626874	0.1174296	1.39	0.168
α_0	0.000298	0.0003877	0.77	0.443
Panel E. Estimates from the AEC Marginal Model				
ϑ_1	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_1	0.1391981	0.1193005	1.17	0.245
α_0	-0.0002894	0.0006605	-0.44	0.662

Notes: *SIC indicates one lag should be included

Table S3. Zivot and Andrews Test for Structural Breaks

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

Source: Authors' elaboration

Table S4. Further Cointegration Tests Restricting a Trend and a Structural Break to the Cointegration 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

Notes: Two lags included (based on the Schwartz information criterion).
Source: Authors' elaboration

Table S5. Granger Causality Tests on VAR(1) and VAR(2)

Equation	Excluded	χ^2	df	p-value
VAR(2)				
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 S6. Demand and Supply Shifters

Label	Short Description	Source	Frequency	Unit-Root[a]
$\mathbf{Z_r^1}$	Unemployment Rate − Seasonally Adjusted	Eurostat	Monthly	Yes
$\mathbf{Z_r^2}$	GDP and main components (output, expenditure and income)	Eurostat	Quarterly	Yes
$\mathbf{Z_p^1}$	Industrial Inputs Price	IMF	Monthly	Yes
$\mathbf{Z_p^2}$	Energy Price Index	IMF	Monthly	Yes
$\mathbf{P_{EU}}$	Weighted Average EU Milk Price	MMO (EC)	Monthly	Yes

Notes: [a]We tested for the presence of unit root by ADF-GLS, PP, and KPSS tests. Results are available upon request.

Table S7. ECM Model Results for Demand and Supply Shifters

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

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

Table S8. Model Results for the ASRM with 1 and 2 Lags, and the AECMSR with Two Lags and Asymmetries

	ASRM 1		ASRM 2		AECMSR 3	
	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

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