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E-mail addresses: Octavio.Fernandez-Amador@ubik.ac.at, Josef.Baumgartner@wifo.ac.at, jesus.crespocuaresma@wu.ac.at

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Medieninhaber (Verleger), Hersteller: Österreichisches Institut für Wirtschaftsforschung • 1030 Wien, Arsenal, Objekt 20 • Tel. (43 1) 798 26 01-0 • Fax (43 1) 798 93 86 • http://www.wifo.ac.at/ • Verlags- und Herstellungsort: Wien

Milking The Prices: The Role of Asymmetries in the Price Transmission Mechanism for Milk Products in Austria*

Octavio Fernández-Amador† Josef Baumgartner‡

Jesús Crespo-Cuaresma§

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Abstract

We assess empirically the vertical price transmission mechanism between producer and consumer prices of milk products in Austria using monthly data for the period from January 1996 to February 2010. We consider explicitly the existence of asymmetries in the adjustment to the long-run equilibrium using two different types of threshold vector error correction (VEC) models, where an inaction band in the adjustment to the long-run relationship is defined and alternatively where price dynamics differ between periods of increasing and decreasing trends in causal prices. Our results indicate that asymmetries play an important role in the pass-through of prices for milk products in Austria. We provide statistical evidence concerning the fact that the adjustment only tends to take place when deviations from the equilibrium are large enough. Milk, dairy and cheese products and butter tend to remain in positive margins (measured as deviations from the long-run equilibrium) for the retailers' side. The explicit modeling of nonlinearities does not improves out-of-sample forecasting performance.

Keywords: Asymmetric price transmission, threshold models, cointegration, milk prices.

JEL classification: C32, L11, Q13.

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†Corresponding author: University of Innsbruck, SOWI Gebäude, Department of Economics, Universitätsstrasse 15, 6020 Innsbruck (Austria). E-mail address: octavio.fernandez-amador@uibk.ac.at.

‡Austrian Institute of Economic Research (WIFO), Arsenal Object 20, A-1030 Vienna (Austria). E-mail address: Josef.Baumgartner@wifo.ac.at.

§Vienna University of Economics and Business, Department of Economics, Institute for Monetary and Public Policy, Augasse 2-6 1090 Vienna (Austria), International Institute of Applied Systems Analysis (IIASA) and Austrian Institute of Economic Research (WIFO). E-mail address: jesus.crespocuaresma@wu.ac.at.

1 Introduction

Contradicting standard economic theory, empirical research has usually pointed out the existence of asymmetric price transmission (hereafter, APT) from input to output prices (see for example, Geweke, 2004, Meyer and von Cramon-Taubadel, 2004, or Frey and Manera, 2007, for surveys considering gasoline and agricultural markets). In particular, empirical results tend to support larger consumer prices reactions to increases in cost prices than in the case of decreases, both in terms of speed and magnitude of the adjustment (Peltzman, 2000).

Several theoretical underpinnings have been proposed to explain why the pass-through between input and output prices tends to be neither instantaneous nor symmetric. The literature has analyzed some characteristics of the markets related to menu costs at the retail level (Azzam, 1999), perishability of products (Ward, 1982) and storing systems at retail and production levels (Reagan and Weitzman, 1982), search costs in local markets (Benson and Faminow, 1985), public intervention (Kinnucan and Forker, 1987), and market power at retail level (Peltzman, 2000). McCorriston *et al.* (2001) show how in an equilibrium displacement model in the spirit of Gardner (1975), allowing for market power in the food industry under the assumption of non-constant returns to scale, market power is not necessarily a cause for price asymmetries in markets at presence of increasing returns to scale. Using a similar argumentation, but in the framework of horizontal price transmission, Azzam (1999) concludes that asymmetries cannot be systematically attributed to a lack of competition, existing even in a competitive framework. Weldegebriel (2004) and Lloyd *et al.* (2009) point to the impact of the combination of oligopoly and oligopsony (buyer) power in a multi-stage vertical price transmission model. Furthermore, Xia (2009) highlights the importance of the functional forms of farm supply and retail demand on the outcome of (possibly imperfect) price transmission.

The modern empirical literature on price transmission usually adopts a non-structural approach for modeling APT. Within this approach, the relationship among non-stationary series of prices is modeled within the skeleton of a vector error correction (VEC) model, where it is possible to cast APT and other types of non-linearities in either the long run relationship or the shorter run adjustment dynamics. A useful class of models parametrizes the asymmetry by assuming a threshold adjustment to the cointegration relationship. Threshold VEC specifications allow us to model asymmetries which may imply, given the proper structure to the model, the existence of an inaction band of price combinations in which there is no response to deviations from the long run relationship (due to, for instance, fixed costs), or a different adjustment depending on the sign of the change in the defined causal price, allowing to assess statistically the hypothesis of an asymmetric impact of causal price changes in the caused prices. This asymmetry needs not take place immediately, but after a period in which the relevant information is assimilated by market agents.

The empirical literature remarks the existence of APT in food markets and usually the chain of causality goes from downstream to upstream, independently of the subsector and country analyzed. In order to characterize price transmission among retail, wholesale, and shipping-point prices for a subset of fresh vegetables in US, Ward (1982) makes use of Wolfram's (1971) asymmetry modeling procedure and shows that wholesale price changes are not totally reflected at the retail level, whereas the later adjusts to decreases in wholesale price. In contrast, the adjustment of shipping point prices is more fully when wholesale price decreases than when it increases. Ward (1982) suggests oligopolistic structure at the retail level, perishability of products and the possibility of reducing sales as the potential major explanation underlying the asymmetric pass-through found. Analyzing the US beef sector for the period 1981 to 1998, Goodwin and Holt (1999) find support for the general assumption that the line of causality on price transmission is from farm to wholesale and to retail level, though also from wholesale to farm level. Nevertheless, their findings point at the fact that asymmetries are modest and could even be of no economic relevance. Using frequency domain regressions, Miller and Hayenga (2001) analyze the pork meat sector in interior Iowa-Southern Minnesota for the period 1981 to 1995 and find support for retail price asymmetric transmission for low-frequency dynamics of wholesale prices, while in the relationship between farm and wholesale prices no APT is found at any frequency. Abdulai (2002) also offers evidence of asymmetries in pricing behavior of retailers in the pork sector in Switzerland during the period 1988 to 1997. The causality line appears to be from producer to retail levels and, in particular, increases in the producer price that induce a reduction in the margin of the retailers are passed on to retail level faster than reductions in the producer price that imply an increase in the marketing margin. Ben-Kaabia and Gil (2007) find evidence for full price transmission in the long-run in the Spanish lamb sector for the period 1996 to 2002. However, their results show asymmetric adjustments and benefits for the retailers independently of the size, sign or origin of the price shocks. Also, Vavra and Goodwin (2005) find significant asymmetries in the farm, wholesale and retail chain for US beef, chicken and eggs sectors.

Concerning dairy products, the empirical literature has shown similar results. Kinnucan and Forker's (1987) results highlight that asymmetries in both magnitude and time of response are found in retail prices of dairy products (fluid milk, cheese, butter, and ice cream) in the US, with larger and speedier reactions when farm prices are increasing. Serra and Goodwin (2003) find evidence for APT in dairy products in Spain. However, these asymmetries do not seem to be present in highly perishable dairy products. In accordance to McCorriston *et al.* (2001), their results do not suggest a relationship between APT and market concentration. Based on a dynamic reduced-form model of APT, Chavas and Mehta (2004) analyze the butter market in the US for the period 1980 to 2001. They find strong support for asymmetry in the adjustment of retail prices, with a stronger reaction when confronting wholesale price increases than when wholesale price decreases. However, the evidence of APT for wholesale adjustments is weak and based on the asymmetry of retail price adjustments. These authors suggest search costs, menu costs and imperfect competition as causes of the asymmetry at the retail level.

A broader perspective is taken by Peltzman (2000) and Gwin (2009). Peltzman (2000) analyzes 15 2-digit Standard Industrial Classification (SIC) subsectors in US where a single input is the major cost component and finds evidence on APT as a stylized fact. APT appears as a characteristic of competitive and oligopolistic market structures and thus, no evidence on a market concentration foundation for asymmetries is obtained. Also, inventory holdings and menu costs are rejected as plausible explanations for APT. Even though adjustment costs would be a reasonable explanation to his results, Peltzman finds that less input volatility and fragmentation of the supply chain are behind the existence of asymmetries. Gwin (2009) analyzes 269 industries over the 24 2-digit sectors of the North American Industrial Classification System (NAICS) during the period 1966-2006. The results suggest that there is little support for the existence of a economy-wide APT mechanism, finding evidence on APT in nondurable goods (food) and natural resource manufacturing, but not in mining, durable goods manufacturing and service sectors. Due to the differences in price transmission among sectors, Gwin highlights inventory management as the potential explanation for asymmetries.

In this paper we assess the existence of relevant asymmetries in milk products markets in Austria in terms of both speed and magnitude of the changes in consumer and producer prices. We evaluate the vertical price transmission mechanism between consumer and farm-gate producer prices for monthly data of milk, dairy products, cheese and butter for the period 1996 to 2010. We explicitly model potential asymmetries in the transmission mechanism using (a) threshold VEC (TVEC) models, which define an inaction band in the adjustment to the long-run relationship and (b) models where price dynamics differ between periods of increasing and decreasing trends in causal prices (we dub this type of specification SIGN model). We also evaluate the out-of-sample forecasting ability of the estimated asymmetric models and compare them to their linear counterparts.

Our findings give evidence of asymmetries in retail and producer price relationships which materialize in the existence of a band of inaction around the cointegration relationships which link the prices of milk and dairy products in the long run. These asymmetries show persistence in the regime above the inaction band and quick reversion when below the band for all consumer products considered. When modeling the asymmetry in terms of the trend of causal prices, the trigger appears to take place over relatively long periods (about 1 year) for dairy and cheese products whereas for milk and butter it takes just one month. Impulse response functions reflect the trend of the markets to establish the actual relationship of retail and producers prices beyond the band of inaction around the long-run relationship. Our out-of-sample forecasting exercise shows that the estimated models possess good predicting abilities for consumer prices, but do not out-perform linear models.

This paper is structured as follows. Section 2 and 3 describe the main characteristics of the Austrian milk sector and the data, respectively. Section 4 presents the results of linear VEC models. We present the non-linear approach in Section 5, assessing the pass-through in prices by using two types of asymmetric models and analyzing the impulse response functions of such models. Section 6 evaluates the out-of-sample forecasting properties of the proposed models and Section 7 concludes.

2 The main characteristics of the Austrian milk and dairy sector

The raw milk production system in the European Union (EU) is still highly regulated. In 1984 a strict milk quota system, and a reference price for raw milk and intervention prices for butter and skim milk powder were introduced. Hence, the quantity and prices were regulated to reduce overproduction. From 1988 onwards several reforms in this regulatory framework were introduced to allow market forces to play an increasing role.[1] However, the quantity of supply of raw milk is still limited by the quota system, though prices are much less so.[2] Consequently, price changes have become more and more a signal of changes in demand.

As a consequence of the accession to the EU in January 1995, Austria had to implement the EU common market legal framework, which introduced a new set of regulations and stronger international competition in particular to the former highly sheltered agricultural and food processing sector. As a result, the Austrian milk-producing and dairy sector experienced significant changes in the last 15 years resulting in a decrease in the number of farms and in the cowherd, and increasing concentration in the milk-processing industry and among food retailers.

Milk farmers in Austria are of a small scale compared to EU-15 (10 cows per farm in Austria as opposed to 35 in the EU-15) and typically located in alpine regions (around 65% of raw milk production in Austria compared to 12% in the EU-15), where production conditions are tough and alternative production possibilities are rather limited. Consequently, the average annual yield per dairy cow is 10% lower than in the EU-15 (and 28% lower than in the most productive European countries, Denmark and Sweden).[3]

From 1995 to 2008 the number of milk producers in Austria decreased from around 77,000 to 42,000 and the herd of dairy cows shrank from 638,000 to 527,000, whereas the total volume of production increased from 2.9 to 3.2 million tons. An atomized raw milk production sector is confronted with a quite concentrated dairy sector. In 2008 the Austrian top three dairy companies had a market share of almost 55%. However, the next stage in the production chain is even more concentrated. With a market share of 78.5% of the top three food retailers in Austria in 2008, this sector is among the most concentrated in Europe.[4]

3 Data description and time series properties

We analyze publicly available monthly time series for agricultural producer prices of raw milk and consumer price data for milk, dairy products, cheese and butter based on the Austrian consumer price index (CPI) items obtained by Statistics Austria. Taking into account the regime change introduced by the accession to the EU in January 1995 and a major change in the construction of CPI basket operative from January 1996 onwards, the sample period for the empirical analysis spans from January 1996 to February 2010. Statistics Austria provides agricultural producer prices for milk in value terms for two types of raw

[1]The main initiatives were the McSherry reform in 1992, the Agenda 2000, the Common Agricultural Policy reform in 2003, and the Health Check reform in 2008. To prevent milk farmers from income losses as result from cuts in intervention prices they were compensated by direct payments. See Oskam *et al.* (2010) for further details.

[2]During the period 1990 to 2008 the total annual raw milk production in the EU-15 was in the range from 118.8 to 122.9 million tons.

[3]Figures are for 2005 and from Eurostat and Kirner *et al.* (2007).

[4]Figures are from the annual report of the Austrian agricultural sector (Austrian Minstry of Agriculture, 2009) and from Nielsen (2009).

milk, depending on the content of fat. We use the arithmetic mean of these two series as the farm-gate producer price for raw milk. In the goods and services basket for the Austrian CPI 12 single items are included for milk, dairy products, cheese and butter, representing a weight of 1.5% of an average households expenditure in 1995. For dairy products (7 items, milkshake, sour cream, whipped cream, curd cheese, condensed milk, fruit flavored yogurt and curd cream with fruits) and for cheese (3 items, Emmental cheese, Gouda cheese and Camembert cheese) a composite price index is generated as a weighted average from the single items, respectively. In order to make the producer price series comparable with retail prices, index numbers (1996M1=100) were constructed. For the analysis we take natural logarithms of all series and seasonally adjust them applying the TRAMO/SEATS method (Gómez and Maravall, 1996). Due to the fact that the results for dairy and cheese categories are very similar to those of the respective composite indices, we concentrate on the results of the corresponding indices (denoted mi_{cp}, da_{cp}, ch_{cp} and bu_{cp} for milk, dairy products, cheese and butter, respectively).[5]

Augmented Dickey and Fuller (1979, ADF) and Kwiatkowski et al. (1992, KPSS) tests were carried out in order to assess the order of integration of the series in the models. Both tests give evidence of the existence of a unit root in the series under consideration, with the only exception of milk producer prices. For this price index, the results are contradictory depending on the test statistic and setting used. We consider the series to be integrated of order one and use the framework of error correction models in the presence of cointegrating relationships for the bivariate modeling exercise.

Table 1: **Unit root test results**

	Setting with intercept		Setting with intercept and linear trend	
	ADF test stat.	KPSS test stat.	ADF test stat.	KPSS test stat.
$\ln(mi_{cp})$	0.944	1.446***	-1.952	0.187**
$\ln(da_{cp})$	-0.390	1.354***	-2.845	0.190**
$\ln(ch_{cp})$	-0.457	1.323***	-2.286	0.265***
$\ln(bu_{cp})$	-2.094	1.037***	-2.846	0.086
$\ln(mi_{pp})$	-3.117**	0.433*	-4.055***	0.057

Note: *, ** and *** stands for significance at the 10, 5 and 1% level, respectively.

4 Price transmission of milk prices in a linear framework

In a first step we analyze the price transmission of milk product prices using a linear VEC representation of the bivariate dynamics of producer and consumer prices. We thus assume that the price dynamics can be represented as

$$\Delta p_t = \gamma_0 + \alpha\beta' p_{t-1} + \sum_{j=1}^{l} \Gamma_j \Delta p_{t-j} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \Sigma) \tag{1}$$

[5] Additional information on all 12 individual products are available from the authors upon request.

where $p_t = (p_t^c, p_t^p)'$ is a vector composed by the consumer price of a given product and the corresponding producer price (both in logs), γ_0 is a vector of parameters, $\beta' p_{t-1}$ defines the long-run equilibrium relationship between the two price levels, given by the cointegrating vector $\beta' = (1\ -\beta_1)$ and $\alpha = (\alpha_1\ \alpha_2)'$ is a vector of adjustment parameters. Price dynamics also depend on previous changes in both variables up to the l-th lag through the parameter matrices Γ_j for $j = 1, \ldots, l$.

We estimate the VEC model for two possible specifications of the long-run relationship. On the one hand, we obtain estimates of the long-run elasticity by estimating the system given by (1), including the long run relationship, using maximum likelihood methods. On the other hand, we restrict the long-run elasticity to be equal to unity ($\beta_1 = 1$), as economic theory would suggest in a constant mark-up framework. The optimal lag length in (1) is obtained by minimizing the Schwarz (1978) information criterion (Bayesian Information Criterion, BIC) over lag lengths ranging from one to eighteen.

Johansen's (1991) cointegration test is performed for all the pairs considered in the bivariate specification for unrestricted and restricted ($\beta_1 = 1$) long-run specifications. Model selection using BIC resulted in specifications with two lags in the first differences of the price vector for all cases. The results of the estimation of the linear models are presented in Table 2 and give strong evidence of the existence of cointegration between the pairs of variables considered, for unrestricted specifications for milk, dairy and cheese indices, and restricted specifications, in the case of butter. In order to assess the stylized facts of the lag structure of the relationship, we also performed Granger's (1969) causality tests on the vector autoregressive model in first differences defined by specification (1) without the error correction term. As shown in Table 2, the direction of the relationship is from producer prices to the consumer prices at 5% of significance in all the price pairs considered.

Table 2 also presents the estimates of the long-run elasticity and the adjustment parameters for the linear VEC models given by (1) for unrestricted and unit-elasticity specifications. The adjustment to the long-run attractor takes place through changes in the producer price in all models. For milk and cheese there is also weak evidence that some adjustment (but with an unexpected sign in the case of cheese) is carried out through changes from consumer prices as well. The long-run elasticities of milk, cheese and dairy products reflect a more than proportional pass-through between prices.

5 Modeling asymmetric price transmission

We consider two types of asymmetric adjustment models for the system formed by producer and consumer prices. We firstly consider a model where the adjustment to the cointegration relationship takes place exclusively for relatively large deviation of the equilibrium relationship, while a band of inaction appears for smaller deviations from the long-run relationship. We also consider models where the asymmetry is triggered by the short-run trend in causal prices, so that the parameters of the error correction model (but not the long-run elasticities) differ between periods which follow increasing causal prices and those preceded by decreasing causal prices.

Table 2: **Linear VEC models**

$\ln(mi_{cp}) - \ln(mi_{pp})$	Unrestricted	Restricted
Long-run elast.:	2.7242*** (0.7132)	1
EC Adjustment: $\ln(mi_{cp})$	0.0066* (0.0039)	0.0072 (0.0096)
EC Adjustment: $\ln(mi_{pp})$	0.0183*** (0.0066)	0.0309* (0.0163)
Lag length:	2	2
P-val coint. test ($H_0 : 1$ CEs)	0.6095	
Unit long-run elasticity test (p-value)		0.021104
Causality test	F-Stat.	P-value
H_0 : No $\ln(mi_{pp}) \rightarrow \ln(mi_{cp})$	4.06781	0.0189
H_0 : No $\ln(mi_{cp}) \rightarrow \ln(mi_{pp})$	2.24412	0.1093
$\ln(da_{cp}) - \ln(mi_{pp})$	Unrestricted	Restricted
Long-run elast.:	3.4061*** (0.8484)	1
EC Adjustment: $\ln(da_cp)$	-0.0021 (0.0019)	-0.0076 (0.0075)
EC Adjustment: $\ln(mi_pp)$	0.0148*** (0.0049)	0.0458** (0.0196)
Lag length:	2	2
P-val coint. test ($H_0 : 1$ CEs)	0.8101	
Unit long-run elasticity test (p-value)		0.036532
Causality test	F-Stat.	P-value
H_0 : No $\ln(mi_{pp}) \rightarrow \ln(da_{cp})$	8.13299	0.0004
H_0 : No $\ln(da_{cp}) \rightarrow \ln(mi_{pp})$	2.15983	0.1187
$\ln(ch_{cp}) - \ln(mi_{pp})$	Unrestricted	Restricted
Long-run elast.:	5.7261*** (1.3024)	1
EC Adjustment: $\ln(ch_{cp})$	-0.0033* (0.0018)	-0.0112 (0.0088)
EC Adjustment: $\ln(mi_{pp})$	0.0091*** (0.0030)	0.0138 (0.0147)
Lag length:	2	2
P-val coint. test ($H_0 : 1$ CEs)	0.5221	
Unit long-run elasticity test (p-value)		0.000514
Causality test	F-Stat.	P-value
H_0 : No $\ln(mi_{pp}) \rightarrow \ln(ch_{cp})$	4.33201	0.0147
H_0 : No $\ln(ch_{cp}) \rightarrow \ln(mi_{pp})$	1.31275	0.2719
$\ln(bu_{cp}) - \ln(mi_{pp})$	Unrestricted	Restricted
Long-run elast.:	0.9855*** (0.1713)	1
EC Adjustment: $\ln(bu_{cp})$	-0.0142 (0.0234)	-0.0129 (0.0230)
EC Adjustment: $\ln(mi_{pp})$	0.0765*** (0.0260)	0.0757*** (0.0256)
Lag length:	2	2
P-val coint. test ($H_0 : 1$ CEs)	0.0188	
Unit long-run elasticity test (p-value)		0.964649
Causality test	F-Stat.	P-value
H_0 : No $\ln(mi_{pp}) \rightarrow \ln(bu_{cp})$	6.72453	0.0016
H_0 : No $\ln(bu_{cp}) \rightarrow \ln(mi_{pp})$	0.55824	0.5733

Note: *, ** and *** stands for significance at the 10, 5 and 1% level, respectively.

Our TVEC model postulates potentially different adjustment parameters above and below the band of inaction, and no adjustment to the long-run equilibrium within the band. It is thus parametrized as follows,

$$\Delta p_t = \begin{cases} \gamma_{0L} + \alpha_L \beta' p_{t-1} + \sum_{j=1}^{l} \Gamma_{Lj} \Delta p_{t-j} + \xi_{Lt}, & \text{if } \beta' p_{t-1} \leq \theta_L, \\ \gamma_{0M} + \sum_{j=1}^{l} \Gamma_{Mj} \Delta p_{t-j} + \xi_{Mt}, & \text{if } \theta_L < \beta' p_{t-1} < \theta_H, \\ \gamma_{0H} + \alpha_H \beta' p_{t-1} + \sum_{j=1}^{l} \Gamma_{Hj} \Delta p_{t-j} + \xi_{Ht}, & \text{if } \beta' p_{t-1} \geq \theta_H, \end{cases} \tag{2}$$

where the band of inaction is thus given by the interval $(\theta_L,\ \theta_H)$, defined in the range of deviations from the cointegration relationship. Based on the results of the linear model and in order to simplify the interpretation of the results we fix the cointegrating vector $\beta = (1\ \ -\beta_1)$ to both cases considered in the linear specification, where β_1 is the estimated parameter from the linear VEC estimation in the unrestricted case and one in the restricted unitary elasticity case. With the same aim, we restrict the lag length of the TVEC model to that of the linear one. The model in (2) in the restricted case, for instance, hypothesizes that price adjustment to the long-run equilibrium only takes place if the price difference between consumer and producer prices exceeds θ_H or falls below θ_L and the speed of such an adjustment is potentially different in both regimes. Furthermore, intercepts and parameters defining the short-run dynamics in the regimes defined by these thresholds are also allowed to differ across regimes. In practice, we do not set the thresholds exogenously, but estimate them as

$$(\widehat{\theta}_L\ \widehat{\theta}_H) = \arg\min_{\theta_L \theta_H}\ \mathrm{SSR}(\theta_L\ \theta_H), \tag{3}$$

where $\mathrm{SSR}(\theta_L\ \theta_H)$ is the sum of squared residuals of the model with thresholds given by θ_L and θ_H. The sum of squared residuals is minimized using a grid search over $(\theta_L\ \theta_H)$. We design the grid search so that at least 10% of the observations fall in the medium regime and 15% in the extreme regimes, in order to avoid model estimates based on regimes with too few observations. The estimated thresholds $(\hat{\theta}_L,\ \hat{\theta}_H)$ are not restricted to be symmetric around the band of inaction.

In addition, we use an alternative modeling strategy by considering that the pass-through between prices is different depending on the past trend of the causal price for a period which the market takes as the relevant information period. Thus, the APT model based on increases versus decreases of the causal price is given by the following specification (SIGN model),

$$\Delta p_t = \begin{cases} \phi_{0L} + \delta_L \beta' p_{t-1} + \sum_{j=1}^{l} \Phi_{Lj} \Delta p_{t-j} + \upsilon_{Lt}, & \text{if } p_{t-1} - p_{t-w} \leq 0, \\ \phi_{0H} + \delta_H \beta' p_{t-1} + \sum_{j=1}^{l} \Phi_{Hj} \Delta p_{t-j} + \upsilon_{Ht}, & \text{if } p_{t-1} - p_{t-w} > 0, \end{cases} \tag{4}$$

where the adjustment parameter and short-run dynamics parametrized through the Φ matrices differ depending on whether the growth rate of producer prices over the last $\omega - 1$ periods was positive or negative. The parameter ω is estimated using a grid search over a reasonable set of lags. In our case, we search in the set of lag lengths $\omega = 2, \ldots, 13$, which range from considering the change in producer prices in the last month to considering the trend in the last year. Therefore, we estimate the period of information of the market as

$$\hat{\omega} = \arg\min_{\omega} \ \mathrm{SSR}(\omega), \tag{5}$$

where SSR($\cdot$) is the sum of squared residuals, minimized now using a grid search over (ω). We design the grid search to ensure that at least 15% of the observations fall in one regime, in order to avoid model estimates based on regimes with too few observations.

The main results of the estimation of nonlinear models are presented in Table 3 and the deviations from the cointegration relationships together with the estimated thresholds for the TVEC specifications are presented in Figure 1. Table 3 presents also the results of a nonlinearity test in the spirit of Hansen (1996). It is well known that the test of a model such as (1) against the nonlinear alternative (2) suffers from the problem that the threshold parameters are nuisance parameters which are not identified under the null hypothesis of linearity. This implies that the usual likelihood ratio (LR) test for model (2) against model (1) cannot be evaluated using standard probability distributions (see Andrews and Ploberger, 1994). We obtain a simulated distribution of the test statistic under the null of linearity as follows. For a given pair of prices, we use the estimates of the parameters in model (1) together with simulated shocks in order to obtain price paths under the maintained linearity assumption. We estimate nonlinear models such as (2) for each of the simulated datasets and calculate the corresponding LR test statistic. By repeating this procedure 1,000 times, we reconstruct the distribution of the test statistic under the null hypothesis. The corresponding p-value is given by the proportion of simulated test statistics which exceed the value obtained using the actual data. LR tests were carried out for the null of linearity against the SIGN model of (4) and for the null of (5) against the alternative of the TVEC model specified in (2).

In every bivariate model considered the TVEC model appears as the preferred one when comparing it to the linear VEC and the SIGN model. Therefore, consumer and producer prices, if they adjust to equilibrium, they only do it as a response to relatively large deviations from long-run equilibrium. However, as shown in Figure 1 the widths of the bands of inaction given by the estimated thresholds differ considerably, from a very narrow one for butter to the widest band for cheese. The ranges of the deviations from the long-run equilibria differ considerably, too. Cheese shows higher volatility, whereas milk and dairy products display similar figures, and butter shows the lowest range. However, their dynamics look like very similar. Furthermore, the adjustment to equilibrium seems to follow an asymmetric pattern. For all products considered in our analysis, there is a tendency of a (very) slow adjustment if deviations from long-run equilibria are positive, whereas negative deviations are corrected much faster. This stylized fact is interpreted as a (small) positive margin which benefits retailers. For milk none of the adjustment coefficients is significant. For dairy products and butter the adjustment comes from the consumer price from below and from the producer price from above. The adjustment coefficients are the largest for butter, yielding the smallest margin. Cheese products show a similar adjustment pattern as dairy products. The adjustment takes also place from above for consumer prices, but with an unexpected negative sign.

When modeling the cointegration relationship depending on the trend of the causal variable determined, the producer price, the estimate of the lag in the price change that triggers the

Table 3: **Threshold VEC models**

$\ln(mi_{cp}) - \ln(mi_{pp})$	TVEC Unrestricted	SIGN Unrestricted
Long-run elast.:	2.7242	2.7242
EC Adjustment (above/increasing trend): $\ln(mi_{cp})$	-0.0029 (0.0100)	0.0020 (0.0061)
EC Adjustment (below/decreasing trend): $\ln(mi_{cp})$	0.0086 (0.0130)	0.0139** (0.0060)
EC Adjustment (above/increasing trend): $\ln(mi_{pp})$	0.0278 (0.0170)	0.0046 (0.0091)
EC Adjustment (below/decreasing trend): $\ln(mi_{pp})$	-0.0183 (0.0220)	0.0394*** (0.0089)
Low threshold:	-0.131485	
Upper threshold:	-0.003001	
Threshold lag length $(\omega - 1)$:		1
LR test (H_0 : Linear model) :	106.0097 (0.0000)	57.7301 (0.0000)
LR test (H_0 : SIGN model) :		48.2797 (0.0000)
$\ln(da_{cp}) - \ln(mi_{pp})$	TVEC Unrestricted	SIGN Unrestricted
Long-run elast.:	3.4061	3.4061
EC Adjustment (above/increasing trend): $\ln(da_{cp})$	-0.0048 (0.0064)	-0.0023 (0.0022)
EC Adjustment (below/decreasing trend): $\ln(da_{cp})$	0.0115** (0.0049)	0.0078 (0.0052)
EC Adjustment (above/increasing trend): $\ln(mi_{pp})$	0.0374** (0.0170)	0.0117** (0.0055)
EC Adjustment (below/decreasing trend): $\ln(mi_{pp})$	0.0187 (0.0130)	0.0742*** (0.0130)
Low threshold:	-0.17363	
Upper threshold:	0.015916	
Threshold lag length $(\omega - 1)$:		12
LR test (H_0 : Linear model) :	119.3508 (0.0000)	64.7407 (0.0000)
LR test (H_0 : SIGN model) :		62.3543 (0.0000)
$\ln(ch_{cp}) - \ln(mi_{pp})$	TVEC Unrestricted	SIGN Unrestricted
Long-run elast.:	5.7261	5.7261
EC Adjustment (above/increasing trend): $\ln(ch_{cp})$	-0.0196*** (0.0068)	-0.0019 (0.0021)
EC Adjustment (below/decreasing trend): $\ln(ch_{cp})$	0.0133*** (0.0047)	-0.0064 (0.0052)
EC Adjustment (above/increasing trend): $\ln(mi_{pp})$	0.0190* (0.0111	0.0064** (0.0033)
EC Adjustment (below/decreasing trend): $\ln(mi_{pp})$	0.0092 (0.0077)	0.0478*** (0.0079)
Low threshold:	-0.301596	
Upper threshold:	0.062167	
Threshold lag length $(\omega - 1)$:		12
LR test (H_0 : Linear model) :	102.6462 (0.0000)	63.3686 (0.0000)
LR test (H_0 : SIGN model) :		29.4331 (0.0000)
$\ln(bu_{cp}) - \ln(mi_{pp})$	TVEC Restricted	SIGN Restricted
Long-run elast.:	1	1
EC Adjustment (above/increasing trend): $\ln(bu_{cp})$	-0.0154 (0.0439)	0.0254 (0.0320)
EC Adjustment (below/decreasing trend): $\ln(bu_{cp})$	0.1653* (0.0915)	-0.0284 (0.0337)
EC Adjustment (above/increasing trend): $\ln(mi_{pp})$	0.0854* (0.0515)	0.0271 (0.0337)
EC Adjustment (below/decreasing trend): $\ln(mi_{pp})$	0.0727 (0.1073)	0.1304*** (0.0355)
Low threshold	-0.038843	
Upper threshold	-0.013852	
Threshold lag length $(\omega - 1)$		1
LR test (H_0 : Linear model) :	53.2870 (0.0490)	41.5226 (0.0000)
LR test (H_0 : SIGN model) :		11.7645 (0.0000)

Note: *, ** and *** stands for significance at the 10, 5 and 1% level, respectively.

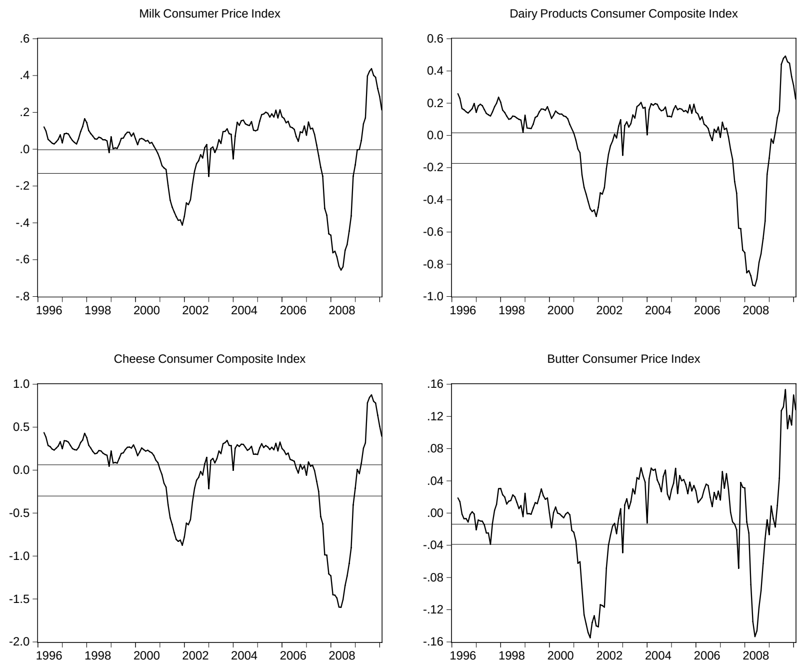

Figure 1: Deviations from cointegration relationships and TVEC thresholds

nonlinearity is one year for dairy and cheese products and one month for milk and butter. Considering the significance of the adjustment, when the causal relationship is from the producer to the consumer price, for all the four product categories the adjustment takes place from the producer side when the trend of change is decreasing, and in the case of dairy and cheese products also when this trend of change is increasing. All of the estimated adjustment parameters have the expected sign. In the case of milk the adjustment takes place also from the side of retailers for a decreasing trend.

The different price dynamics implied by the nonlinear models can be examined by using simulation methods. We performed multiple simulations by starting with a vector of producer and consumer prices in the long-run equilibrium given by the cointegration relationship and applying a shock to the dynamic system that deviates the causal price index of the cointegration relationship by f percentage points. Obtaining a general function which summarizes the dynamics after such a shock is not straight forward in the framework of nonlinear models. On the one hand, the response of a variable to a shock in another one in models such as the TVEC put forward above depends on the history of the variables, in particular on the regime which is active at the moment of the shock. While in the case of linear models differences in the size of the shock do not lead to qualitatively different dynamics in the variables of interest, this is not the case in the nonlinear models considered. Therefore, in spirit of Potter (1995 and 2000) and Koop *et al.* (1996) the generalized impulse response (GIR) functions are estimated as

$$GIR_Y\left(n, \kappa_t, \mu_{t-1}, ..., \mu_{t-j}\right) = E\left[Y_{t+n} | \kappa_t, \mu_{t-1}, ..., \mu_{t-j}\right] - E\left[Y_{t+n} | \mu_{t-1}, ..., \mu_{t-j}\right], \qquad (6)$$

where n are the periods for which the GIR is estimated, ranging from 1 to an horizon h periods (24 months in our case), κ_t is the shock (1% at period $n = 0$), and $\mu_{t-1}, ..., \mu_{t-j}$ is the history of the vector of variables determined on the long-run equilibrium of period $n = 0$ up to lag j, the lag length of the model. For the set of nonlinear models under consideration, the conditional expectations are obtained by simulating the response using a Monte Carlo procedure based on 10,000 replications of the model after the assumed shock takes place.

Figure 2 shows the response of consumer prices over 24 months after a deviation of 1% of the producer price for raw milk from the (sustained) long-run equilibrium computed using the preferred (TVEC) model described above. A detailed analysis of the responses to shocks of different size allows us to draw some general conclusions about the out-of-equilibrium dynamics of the prices under study. First of all, TVEC models reveal that shocks have persistent effects on consumer prices, reflecting the trend for different products to remain in one regime beyond the band of inaction.[6] Secondly, independently of the sign of the shock, the responses of the consumer prices go in the same direction, suggesting that the band of inaction gives a boost to the vector of prices so as for the consumer price to move to the regime above or below the inaction band. Thirdly, these reactions are positive in the case of milk, dairy products and cheese, and negative in the case of butter.

6 Out of sample prediction: Do asymmetries help forecasting?

In a further step, we evaluate whether nonlinear modeling of the price transmission mechanism in prices helps us to improve forecasts of consumer prices. Since some of the regimes in the TVEC and SIGN model are not stable, we concentrate on the accuracy of qualitative forecasts about the direction of change in consumer prices (increases versus decreases). The design of the out-of-sample forecasting exercise is as follows. We estimate the linear and the two asymmetric models using data ranging from January 1996 to January 2004. We use the estimated models to obtain predictions for the direction of changes in consumer prices in the period February 2004 - January 2005. We add the observation corresponding to February 2004 to the sample, reestimate our models and repeat the exercise for the out-of-sample period March 2004 - February 2005. This is repeated until the end of the available sample is

[6] We also simulated shocks of very small size, confirming that this trend to stay in one regime beyond the inaction band origins from the dynamics *within* the band.

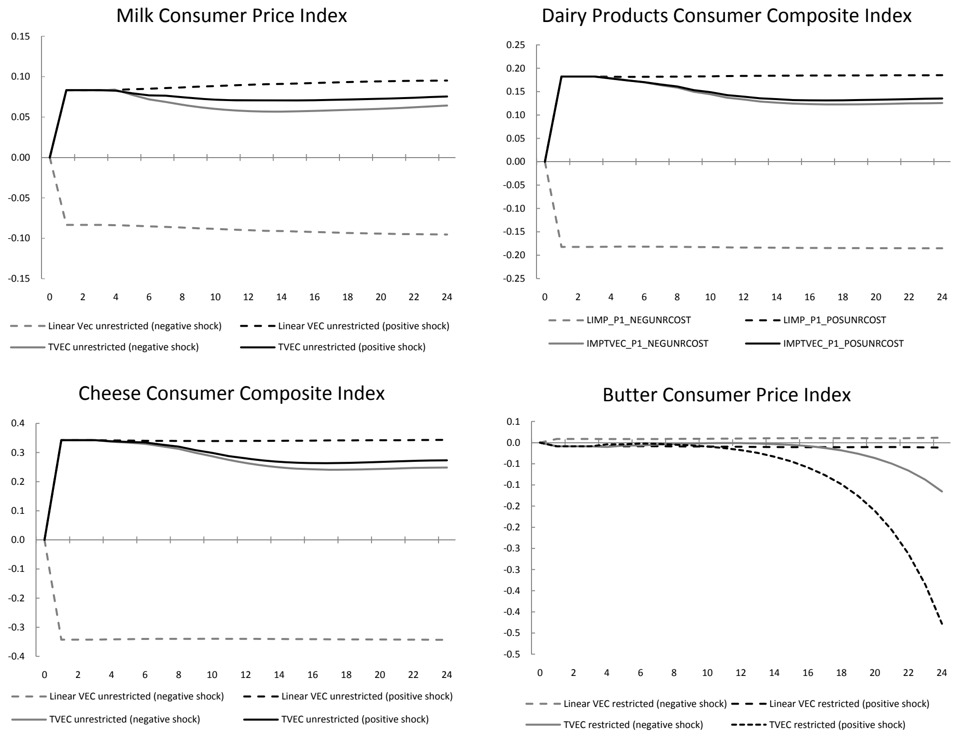

Figure 2: Generalized impulse response functions

reached and the average proportions of correctly forecast directions of change are computed.

The results for 3, 6, 9 and 12 months ahead forecasts for the models considered under the accepted cointegration relationship specification are presented in Figure 3 for consumer prices for milk, dairy and cheese products and butter. In Figure 3 we show the proportion of correct forecasts of direction of change for each forecasting horizon and each model. In general, the forecast performance of non-linear models is good but not superior to linear models. The SIGN model specifications seem to offer a better forecasting performance than the TVEC model. This result is not surprising in the context of results in the forecasting literature which favor parsimonious models, showing the SIGN model as a good combination of parsimony and complexity.

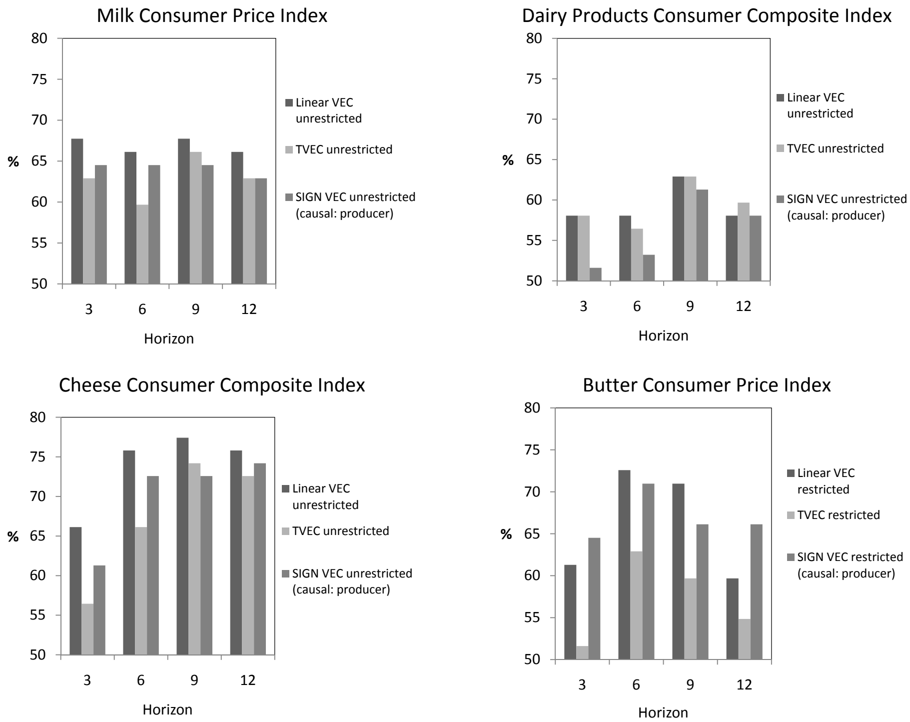

Figure 3: Out of sample forecasting performance

7 Conclusions

In this study we assess the price transmission mechanism between producer and consumer prices of milk and dairy products in Austria using monthly data for the period January 1996 to February 2010. The asymmetric price transmission mechanism between consumer and producer prices for monthly data of milk, dairy and cheese products, and butter is explicitly modeled using threshold VEC models defining an inaction band around the long-run relationship (TVEC models) and models where price dynamics differ between periods of increasing and decreasing trend of change in causal prices (SIGN models). We also analyze the short-run dynamics of the models proposed and their forecast performance.

Our results show robustly that asymmetries play a role in milk and dairy markets in Austria. These asymmetries can be modeled as triggered by the magnitude of the deviation from equilibrium, as well as the trend in prices in a reference period. The preferred models imply that long-run attraction forces seem to be relevant only for relatively large deviations from the equilibrium of the market. For all products persistent positive deviations from the long-run equilibrium are revealed. This situation seems to point to positive mark-ups and benefits for retailers. Impulse response analysis gives further support to the bias of the market when establishing prices beyond the inaction band around the long-run equilibrium. Modeling nonlinearities explicitly does not help to improve the forecast performance.

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