Sergei KHARIN*, Zuzana LAJDOVA and Peter BIELIK****

Price transmission on the Slovak dairy market

There are problems in the functioning of the food supply chain related to price transmission and value-added distribution. Vertical price transmission analysis is an important research area in the aspect of the assessment of impact on the welfare at the producer, processor and retailer levels. The paper investigates vertical price transmission along the whole milk supply chain after the end of European Union milk quotas in the Slovak market using a vector error correction model. Monthly farm-gate, processor and retail prices in the Slovak Republic covering the period from 2010 to 2016 were used in the analysis. Using the Johansen co-integration technique, empirical evidence has been found for two co-integration equations between farm-gate, processor and retail prices. We show that short-term and long-term bilateral causal relationships exist between prices at different market stages. The estimation of the price transmission elasticity supports the assumption that price changes are not transmitted efficiently from one level to another. However, symmetric price transmission exists between farm-gate and processor prices for whole milk in the long term. The perfect price transmission may also be due to recently emerging and strengthening the producer organisations that enable producers support their bargaining position in the supply chain.

Keywords: dairy sector, elasticity, price, vector error correction model

* Voronezh State University of Forestry and Technologies, ulica Timiryazeva 8, 394087 Voronezh, Russia. Corresponding author: kharins03@gmail.com; http://orcid.org/0000-0003-0563-1269
** Slovak University of Agriculture in Nitra, Nitra, Slovakia.

Introduction

Prices drive resource allocation and output mix decisions by economic actors, and price transmission integrates markets vertically and horizontally (Meyer and von Cramon-Traubadel, 2004). As noted by Fousekis *et al*. (2016), vertical price transmission has attracted considerable attention in agricultural economics research for almost 50 years due to the fact that the magnitude and/or the speed at which shocks are transmitted from one market level to another has important welfare and policy implications. Likewise, Goodwin (2006) points out that the degree to which market shocks are transmitted along the marketing chain has long been considered to be an important indicator of the performance of the market.

Bakucs *et al*. (2014) studied explanations for the existence of price (a)symmetries and showed that asymmetric price transmission exists in farm-retail relationships with more fragmented farm structure, higher governmental support and more restrictive regulations on price controls in the retail sector. By contrast, more restrictive regulations on entry barriers in the retail sector and the relative importance of the sector can be favour symmetric farm-retail price transmission. Similarly, Santeramo and von Cramon–Taubadel (2016) mentioned that asymmetric vertical price transmission has been stimulated in several ways such as market power, adjustment costs, inventory management, government interventions, asymmetric information and perishability.

Early analyses typically used simple correlation statistics or ordinary least square regressions to evaluate the links between prices at different markets or processing stages, but these methods have been criticised for not recognising the non-stationary nature of data. Therefore, techniques such as co-integration and error correction models (Akdi and Berument, 2006; Lambert and Miljkovic, 2010; Baek and Koo, 2014; Castillo-Valero and García-Cortijo, 2015; Zhang *et al*., 2017), dealing with non-stationary properties of time series, have been applied since 1987. Recently, nonlinear behaviour in price transmission has been tested using nonlinear threshold techniques (Goodwin and Harper, 2000; Ning and Sun, 2014; Hassouneh *et al*., 2015). The relationship between variables might be locally linear, however globally it exhibits its nonlinear behaviour due to the existence of structural changes in the relationship (Ihle and von Cramon-Taubadel, 2008).

Awokuse and Wang (2009) studied the effect of nonlinear threshold dynamics on asymmetric price transmission for U.S. dairy products (butter, cheese and fluid milk) and confirmed the presence of asymmetric price adjustments for butter and fluid milk, but not for cheese prices. Fałkowski (2010) investigated price transmission between farm and retail levels in Poland by using a vector error correction model (VECM) framework and found that price transmission is influenced by both short- and long-term asymmetries; moreover, the behaviour of prices in the fluid milk sector acts in accordance with the use of market power by the downstream sector. Further evidence of short-term and long-term asymmetries between milk prices of the marketing channel for Poland is provided by Bakucs *et al*. (2012), who concluded that the causality runs from the retail industry to the farm gate and considered, among others, dairy farm structure (individual farms and excessive herd fragmentation in Poland), market structure at the processing level (dairy cooperatives in Poland) and concave spatial demand as causes of (im)perfect pass-through of prices. Similarly, Reziti (2014) used an error correction model to test for asymmetric adjustments in the Greek milk sector and found that retail prices adjust if the producer price increases, not decreases, in the short term. Furthermore, the results confirm asymmetry in the long term, suggesting that retailers exercise market power over producers. Weber *et al*. (2013) show that the time lags in which changes are passed on between the different levels vary and conclude that price asymmetries occur within the supply chain of the German cheese market. In addition, asymmetric threshold VECMs, applied by Serra and Goodwin (2003), reveal asymmetries among farm and retail markets for a variety of dairy products in Spain. The reasons behind the weak response of farm prices to retail price shocks may be partly explained by the lack of an organised contracting system and a scarcity of dairy farmer cooperatives that may limit the market power of farmers relative to the dairy industry, as well as their capacity to negotiate prices.

On the other hand, Weaver and Rosa (2016) provided strong evidence of symmetry in co-movement for the vertical dairy chain in Italy by using a parametric test of asymmetry in a multivariate VECM. Likewise, the price transmission was strong and symmetric for Danish milk from wholesale to retail in the long term (80-85 per cent), surveyed by Jensen and Møller (2007). Additionally, symmetric price transmission was found both in both the long and short terms in Hungary, due to the dominant position of large-scale agricultural enterprises, and FDI in the Hungarian dairy industry and emerged producer organisations; moreover the causality between Hungarian milk prices runs from the farm to the retail sector (Bakucs *et al.*, 2012).

Weldesenbet (2013) demonstrated asymmetric price transmission in the Slovak milk market from 1993 to 2010 in both short and long terms, meaning that retailers and wholesalers react more quickly to producer price increases than to declines. Similar results were obtained by Pokrivcak and Rajcaniova (2014), who stated that the retail sector has strong market power to influence upstream prices. Lajdová and Bielik (2015) used the VECM method to examine price asymmetries for liquid milk (semi-fat and durable semi-fat milk) in the Slovak dairy sector. Their research confirms asymmetric price adjustments and the imperfect market structure with the prevailing power on the demand side.

Milk production in Slovakia decreased significantly during the period 2007-2013 as a consequence of an increase in competitive pressure in the European Union (EU) market, growing imports of milk and milk products to Slovakia, unprofitable production of milk as well as under-capitalised Slovak agriculture (Matošková and Gálik, 2016). Under these circumstances, Slovak raw cow milk producers have suffered significant financial losses. This trend may continue, due to the Russian import ban on EU dairy products and the abolition of the EU milk quota in 2015. The EU market has been flooded by surplus milk and this was followed by a sharp fall in prices. In addition, processors may cancel or not renew existing supply contracts with raw cow milk producers. The past ten years of milk crises caused huge damage to the milk producers: the number of dairy cows fell by almost 31 per cent; milk deliveries declined by almost 15 per cent; the losses of milk producers reached almost EUR 450 million, and almost 35 per cent of enterprises exited milk production (Štefániková, 2017). Even if Slovak agriculture is dominated by large farms, the disproportionate power between small and large farmers who, in the partnership relationship, the mutual distrust between small and large-scale farmers leads to a lack of cooperation or poor cooperation and their weak bargaining power. Moreover, differences in purchase prices (average milk prices in Slovakia do not reach the EU average level, according to Štefániková, 2017) and unequal support mechanisms (the contribution from the national budget the lowest among all surrounding Member States) worsen the competitiveness of the Slovak dairy sector. Retailers can sell imported dairy products at competitive prices, thus the pricing decisions of producers are also driven by contractual relationships between the processors and retailers.

The main aim of this paper is to investigate vertical price transmission along the dairy supply chain in Slovakia in the light of price developments after the abolition of milk quotas in the EU. By focusing on the latest price developments after milk quota abolition, this study seeks to fill a gap in the literature. It also explores how market changes have altered vertical price transmission, and whether asymmetric price transmission still prevails in the supply chain.

Methodology

Econometric time series techniques were adopted for vertical price transmission analysis. The influence of price at one market stage on price at another is investigated using multiple linear regressions. Vertical price transmission analysis follows the algorithm outlined in Table 1. For the whole milk prices (farm-gate, processor and retail), the following steps have been implemented to identify the appropriate econometric model.

To avoid model misspecification, as a preliminary step of our price series analysis, we tested all the variables for the presence of unit root. For this purpose, several methodological options are available including the Augmented Dickey-Fuller [ADF] test (Dickey and Fuller, 1979) and the Phillips-Perron [PP] test (Phillips and Perron, 1988).

As a standard procedure to test the non-stationarity of price series the ADF test uses following regression:

$$P_t = c + \beta_t + \alpha P_{t-1} + \sum_{i=1}^{k} \psi_i \Delta P_{t-i} + \varepsilon_t \tag{1}$$

where P_t is the natural logarithm of the price, c is the intercept and t is the linear time trend.

In order to select the highest number of lags for our test, we applied the common rule suggested by Schwert (1989). The number of the optimum lags in the models is chosen based on the Akaike (1973) information criterion (AIC).

The PP test builds on ADF test. While the ADF test uses a parametric autoregression, a great advantage of the PP test is that it is non-parametric. The main disadvantage of the PP test is that it works well only in large samples. And it also

Table 1: Algorithm for conducting the vertical price transmission analysis.

Step	Test	Result	Action
1	*Stationarity* test of time series for unit root	Stationarity	Perform test for Granger Causality and estimate vector autoregressive[VAR]model with stationary data.
		Non-stationarity	Move to step 2.
2	*Cointegration* test	Exists	Estimate the long- and short-term relationships within the framework of a VECM.
		No	Perform the Granger Causality test and estimate VAR model using logarithmic prices in first differences

Source: based on Kharin (2015)

shares disadvantages of ADF tests: sensitivity to structural breaks, poor small sample power resulting.

There might be a linear combination of same integrated time series that is stationary. Co-integration analysis is used to estimate long-term price relationships between non-stationary and same integrated variables. Given that some price series might be non-stationary, we applied the Johansen approach to determine whether the three series are co-integrated and to identify the number of co-integrating equations by providing likelihood ratio tests based on the trace statistic and maximum eigen value (Johansen, 1988; Johansen and Joselius, 1990). We relied on trace statistic because it tends to have superior power in empirical papers (Lutkepohl *et al*., 2001). Although co-integration implies that causality exists between price series, it does not indicate the direction of the causal relationship.

If the presence of the long-term relationships between variables is detected, then the vector error correction (VEC) model is estimated.

VECM is a restricted vector autoregressive (VAR) model. The VEC modelling can be written by specifying an unrestricted VAR of order *k* as follows:

$$P_t = c + A_1 P_{t-1} + \ldots + A_k P_{t-k} + \gamma_0 Y_t + \gamma_1 Y_{t-1} + \ldots + \gamma_m Y_{t-m} + \upsilon_t \tag{2}$$

where *c* is the intercept, P_t is a (3x1) vector of all endogenous variables defined in the model (natural logarithms of the farm-gate, processor and retail prices); Y_t is a vector, including all exogenous variables; $A_1 \ldots A_k$ and $\gamma_0 \ldots \gamma_m$ - matrices, including the coefficients to be estimated; υ_t - (3x1) vector of i.i.d normal disturbances with zero mean and covariance matrix $\sum$.

The lag length is determined based on the AIC, the Schwartz-Bayesian Information Criterion (BIC; Schwarz, 1978) and Hannan-Quinn Information Criterion (HQIC; Hannan and Quinn, 1979). When all three agree, the selection is clear, but there may be conflicting results. Ivanov and Kilian (2001) suggest that, in the context of VAR models, AIC tends to be more accurate with monthly data, HQIC works better for quarterly data on samples over 120 observations and BIC works fine with any sample size for quarterly data. Having monthly data, we rely on AIC.

Equation 2 can be adjusted in the form of vector autoregressive in differences and error correction components:

$$\Delta P_t = \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \Pi P_{t-1} + \sum_{j=0}^{m} \gamma_j Y_{t-j} + \upsilon_t \tag{3}$$

Equation 3 is obtained from the level VAR (equation 2) by subtracting P_{t-1} from the both sides. Γ_i is the (3x3) matrix of parameters for an *i* order lag process that capture short-term relationships. Π is the (3x3) matrix that represents long-term dynamics, where $\Pi = \alpha\beta'$, α includes the speed of adjustment coefficients to equilibrium (or error correction term, ECT) and β' is the co-integrating vector in the long term. Since the prices are expressed in logarithms for our analysis, the coefficient β is the long-term elasticity of price transmission.

The VECM indicates the direction of causality among prices and allows us to distinguish between 'short-term' and 'long-term' Granger causality. When the variables are co-integrated, then in the short term, deviations from this long-term equilibrium will feed back on the changes in the dependent variable so as to force the movement towards the long-term equilibrium.

The Wald χ^2-tests (or F-tests) of the differenced explanatory variables give us an indication of the short-term causal effects, whereas the long-term causal relationship is implied through the significance or t-test(s) of the lagged ECT, which contains long-term information since it is derived from the long-term co-integrating relationships. The long-term causality can be tested by looking at the significance of the speed of adjustment (α), which is the coefficient of the ECT.

Results

The price transmission analysis was carried out using monthly observations from January 2010 to November 2016 at the farmer, processor and retailer levels in the Slovak Republic. Observations relate to nominal prices for cow whole milk. The data sources are the 'Price indices and average prices in agriculture and forestry' data of the Statistical Office of the Slovak Republic (available online at http://www.statistics.sk/pls/elisw/MetaInfo.explorer?cmd=go&s=1003&sso=3&so=16) and the online database of the Research Institute of Agricultural and Food Economics in Bratislava (www.vuepp.sk). We use the logarithmic transformation of monthly prices measured in EUR per litre (excluding VAT). From an economic point of view, the transformation allows us to interpret the results in percentage change terms and calculate the price elasticity. Analyses between prices commonly use logarithms because, with trending data, the relative error declines through time (Banerjee *et al*., 1993).

The development of whole milk prices at various levels during the period 2010-2016 is shown in Figures 1 and 2. The mean value of farm-gate price of raw cow milk (class I in quality) equals EUR 0.27 per litre, whereas the average value of processor and consumer prices is EUR 0.52 and 0.72 per litre respectively (Table 2). The coefficient of variation is higher for farm-gate price series in comparison with another price series. Processor and retail prices are less dispersed around the mean value. The standard deviation is rather low (Table 2), so prices are close to the mean of our samples.

Using the methodology described above, we started the price series analysis with the unit root tests. Visual examination of the price series graphs suggests that the model for unit root test should contain a constant and a time trend. Price series stationarity was checked with the ADF and PP tests.

Table 2: Descriptive statistics of whole milk prices (EUR per litre), January 2010 – November 2016.

	Farm-gate	Processor	Retail
Mean	0.26759	0.51566	0.72301
Median	0.28	0.52	0.73
Minimum	0.20	0.40	0.63
Maximum	0.30	0.62	0.82
Std. Dev.	0.028737	0.048795	0.055188
Skewness	-0.68724	-0.28163	-0.11008
C.V.	0.10739	0.094626	0.07633
Kurtosis	-0.87501	-0.17711	-1.2161

Data source: Statistical Office of the Slovak Republic

Table 3: Unit root test results.

Logged price variable	Model	Augmented Dickey-Fuller test				Phillips-Perron test			
		Lag	Levels	Lag	First difference	Lag	Levels	Lag	First difference
Farm-gate	Trend & Intercept	3	-2.276	9	-4.689***	3	-2.280	9	-8.633***
	Intercept only	3	-1.984	9	-2.096	3	-1.897	9	-8.293***
Processor	Trend & Intercept	2	-2.439	1	-5.081***	2	-2.452	1	-10.297***
	Intercept only	2	-1.919	1	-5.118***	2	-1.903	1	-10.326***
Retail	Trend & Intercept	4	-0.418	3	-4.475***	4	-0.479	3	-8.640***
	Intercept only	4	-1.777	3	-2.751*	4	-1.453	3	-8.098***

Note: */**/*** null hypothesis of non-stationarity rejected at the 10%, 5% and 1% levels of significance
Source: own calculations

The optimal lag order was determined based on AIC. The null hypothesis is rejected if the critical value is greater than the test statistic (p-value is less than level of significance). The results are summarised in Table 3. The null hypothesis of stationary price series in levels was rejected for all variables. Tests based on first differences show that all the test statistics are significant. Hence, we can conclude that all price variables are integrated of the order one, I (1).

After establishing the order of integration for each variable, we checked whether they are co-integrated. Given non-stationary price variables of the same order, we ran a Johansen co-integration test in order to reveal if the price series are co-integrated and to determine the number of co-integrating equations. The lag length was identified based on the AIC as a result of VAR modelling with constant and a linear trend. The Johansen co-integration technique discovered two co-integrating equations, according to the *trace* and L_{max} test, as the null hypotheses of $r=0$ and $r\leq 1$ (against the alternatives $r>0$ and $r>1$ respectively) are rejected at the 5 per cent significance level, whereas the null of $r=2$ cannot be rejected (Table 4). Hence, the price series are co-integrated and demonstrate long-term relationships within the analysed period. Therefore, we estimated a VECM with two co-integrating relationships.

The co-integration analysis does not identify any information about the causality direction; however, causality is investigated by means of VECM. Co-integration implies causality in at least one direction. This is indicated by the significant α-parameter. Given co-integration between variables, the VECM is estimated (Table 5). The VECM form with unrestricted constant consists of 12 lags order, which was set by AIC in the VAR model, and three endogenous variables. Ljung-Box (1978) and ARCH tests indicate that the VECM is well specified, residuals do not suffer from serial autocorrelation and there is no heteroscedasticity at the 1 per cent or 5 per cent levels of significance. The Doornik-Hansen (2008) test on the residuals was performed to check whether the residuals are normally distributed. The null hypothesis of multivariate normality cannot be rejected at only the 1 per cent of significance level according to the p-value (0.0141) and the residuals are normally distributed, that is desirable.

Theoretically, the VEC model reveals expected signs for explanatory variables in the long-term period. The coefficients in the long-term relationship are long-term elasticities. Each coefficient measures the corresponding magnitude of change in the dependent variable following a percentage change in a particular explanatory variable. Thus, a 1 per cent increase in retail prices leads to a 0.39 per cent and 0.4 per cent increase in farm-gate and processor prices respectively.

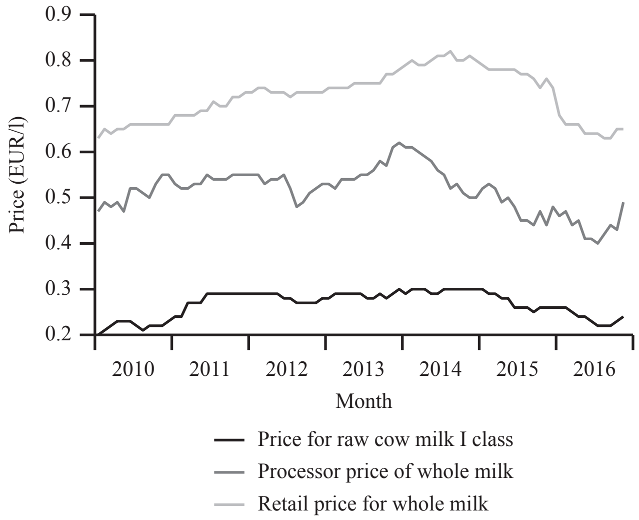

Figure 1: Price series for whole milk in the Slovak Republic, January 2010 – November 2016.
Data source: Statistical Office of the Slovak Republic

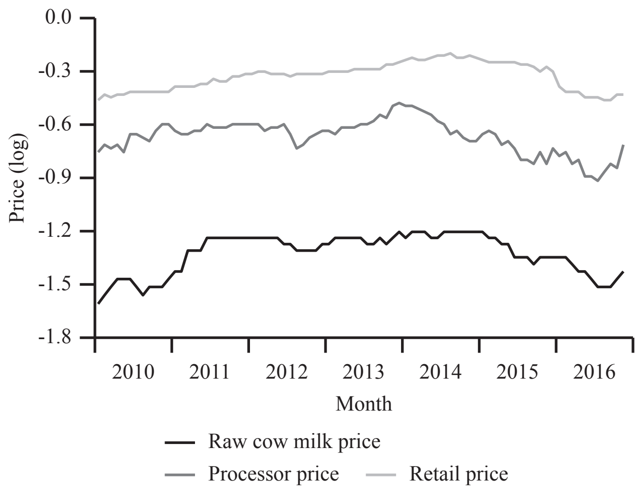

Figure 2: Price series in logarithms for whole milk in the Slovak Republic, January 2010 – November 2016
Data source: Statistical Office of the Slovak Republic

Table 4: Johansen co-integration test.

Hypothesised number of co-integrating equation(s)	Eigen value	Trace test	p-value	Lmax test	p-value
None (r=0)**	0.27284	42.716	0.0008	22.621	0.0284
At most 1 (r≤1)**	0.22887	20.095	0.0083	18.452	0.0087
At most 2 (r≤2)	0.02287	1.6424	0.2000	1.6424	0.2000

Note: ** denotes rejection of the null (0 or 1 co-integration vectors) at the 5% significance level
Source: own calculations

Table 5: Results of VECM estimates.

Co-integrating equation	Model 1		Model 2	
	CointEq1	CointEq2	CointEq1	CointEq2
L_FP_{t-1}	1.0000 (0.0000)	0.0000 (0.0000)	-2.5013 (0.55765)	-1.0006 (0.18824)
L_WP_{t-1}	0.0000 (0.0000)	1.0000 (0.0000)	0.0000 (0.0000)	1.0000 (0.0000)
L_RP_{t-1}	-0.39979 (0.21176)	-0.40003 (0.16784)	1.0000 (0.0000)	0.0000 (0.0000)
Error Correction Term (α)	DL_FP	DL_WP	DL_RP	DL_WP
CointEq1	-0.36287***	0.20625	-0.08511**	0.12774
CointEq2	0.10224	-0.52546**	0.27129**	-0.52546**
Intercept	-0.36987***	-0.03761	0.07415	-0.03761
ΔL_FP_{t-1}	-0.17688	0.30883	0.13875	0.30883
ΔL_FP_{t-2}	-0.06701	-0.10640	-0.05759	-0.10640
ΔL_FP_{t-3}	0.31348**	-0.05179	-0.07421	-0.05179
ΔL_FP_{t-4}	0.18059	-0.39200*	0.05744	-0.39200*
ΔL_FP_{t-5}	0.10377	0.02094	-0.08029	0.02094
ΔL_FP_{t-6}	0.02265	0.43063*	0.01208	0.43063*
ΔL_FP_{t-7}	0.15574	0.48106*	0.11198	0.48106*
ΔL_FP_{t-8}	-0.32855*	-0.74252***	0.18651	-0.74252***
ΔL_FP_{t-9}	-0.02483	-0.06440	0.02716	-0.06440
ΔL_FP_{t-10}	-0.08709	0.00015	-0.04159	0.00015
ΔL_FP_{t-11}	0.10142	0.14376	0.07482	0.14376
ΔL_WP_{t-1}	0.00661	0.44795*	-0.23661*	0.44795*
ΔL_WP_{t-2}	0.07403	0.74568***	-0.19613*	0.74568***
ΔL_WP_{t-3}	0.04870	0.05501	-0.09010	0.05501
ΔL_WP_{t-4}	-0.22968*	0.35608*	-0.15389	0.35608*
ΔL_WP_{t-5}	-0.08944	0.36324*	-0.16923*	0.36324*
ΔL_WP_{t-6}	-0.09985	0.08928	0.03784	0.08928
ΔL_WP_{t-7}	-0.33538**	0.10483	-0.13890	0.10483
ΔL_WP_{t-8}	0.05485	0.41680*	-0.07207	0.41680*
ΔL_WP_{t-9}	-0.00878	0.05248	-0.01379	0.05248
ΔL_WP_{t-10}	-0.10496	-0.08057	-0.12147	-0.08057
ΔL_WP_{t-11}	0.15161	0.60315***	0.00036	0.60315***
ΔL_RP_{t-1}	0.39840	0.91610***	0.06806	0.91610***
ΔL_RP_{t-2}	0.47626**	0.14982	-0.30448*	0.14982
ΔL_RP_{t-3}	0.19113	-0.17617	-0.00286	-0.17617
ΔL_RP_{t-4}	0.42682*	1.05387***	0.04856	1.05387***
ΔL_RP_{t-5}	0.03961	-0.24015	-0.05203	-0.24015
ΔL_RP_{t-6}	0.29301	-0.29509	0.02264	-0.29509
ΔL_RP_{t-7}	0.11392	0.36956	-0.27381	0.36956
ΔL_RP_{t-8}	-0.03783	-0.64339*	-0.20325	-0.64339*
ΔL_RP_{t-9}	0.10722	0.22455	-0.34467**	0.22455
ΔL_RP_{t-10}	-0.10541	0.29304	-0.43781**	0.29304
ΔL_RP_{t-11}	0.50256	0.43628	0.10969	0.43628
R^2	0.75784	0.73669	0.67987	0.73669
Adj R^2	0.51567	0.47338	0.35975	0.47338
F-statistic, *p-value*	3.53e-19	7.97e-24	2.03e-18	7.97e-24
DW-statistic	2.01719	2.10986	2.02855	2.10986
Sum squared residuals	0.01411	0.02701	0.00638	0.02701
S.E. of regression	0.02008	0.02778	0.01351	0.02778
Autocorrelation (Ljung-Box test), *p-value*	0.98	0.351	0.929	0.351
ARCH test, *p-value*	0.8742	0.86146	0.78916	0.86146
Normality of residuals (Doornik-Hansen test), *p-value*	0.0141	0.0141	0.0141	0.0141

Note: */**/*** – statistically significant at the 10%, 5% and 1% levels of significance; standard errors in parentheses; L_FP – farm-gate price in logarithms, L_WP – processor price in logarithms, L_RP – retail price in logarithms
Source: own calculations

In return, a 1 per cent rise of farm-gate price results in an increase in the retail price of 2.5 per cent; therefore, an imperfect market structure is demonstrated, where retailers have a stronger market power than other agents. Interestingly, perfect price transmission exists between farm-gate and processor prices for whole milk. A 1 per cent rise of processor prices leads to an approximately 1 per cent increase in farm-gate prices. The findings also indicate that the ECT coefficients are statistically significant at the 5 per cent level. All the coefficients carry the negative sign, indicating the stability of the system and the convergence towards equilibrium if any disturbance appears in the system. Thus, we can see long-term causality from variable *L_RP* to *L_FP* and vice versa, from *L_RP* to *L_WP* and from *L_FP* to *L_WP*, because the speed of adjustment towards long-term equilibrium is significant and the sign is negative. The ECTs show how fast each variable reaches equilibrium. The higher the value, the faster the reaction. The ECT of ΔL_WP is statistically significant at the 5 per cent level and carries the negative sign. This implies that the restoration to the equilibrium path will not take a long time due to the fact that the α-value (0.52546) is high enough. The ECT of ΔL_FP is statistically significant at the 1 per cent level and carries the negative sign; however, the restoration to the equilibrium path will take longer than the processor price restoration due to the fact that the α-value (0.36287) is smaller. In the case of the retail price movement to equilibrium, it will take rather long time because the α-value (0.08511) is quite small. Thus, the co-integrating vector, in combination with significant and negative error correction terms, indicates long-term causality. The remaining lags in first differences in the VECM are used to test for short-term Granger causality by means of the Wald test. The null of no causality for all the price pairs can be rejected at the 5 per cent level of significance (Table 6). In summary, we found reasonable evidence of short-term causality from the farm-gate to retail prices and vice versa; from processor to farm-gate prices and vice versa; from retail to processor prices and vice versa.

Table 6: Short-term Causality Wald Test results (df = 11).

Dependent variable	Excluded variables	χ^2	p-value
ΔL_FP	ΔL_WP	26.8334	0.00487
	ΔL_RP	34.6938	0.00028
ΔL_WP	ΔL_FP	53.2896	0.00000
	ΔL_RP	54.7952	0.00000
ΔL_RP	ΔL_FP	32.7590	0.00058
	ΔL_WP	22.1784	0.02303

Source: own calculations

Discussion

In this paper, we investigated price transmission along the whole milk supply chain in the Slovak Republic by taking into account the price development after the abolition of milk quotas in the EU. Monthly farm-gate, processor and retail prices in natural logarithms during the period from January 2010 to November 2016 were used in our analysis. Vertical price transmission was evaluated in the co-integration framework, using the Johansen approach, which confirmed the co-integration between price variables and determined two co-integrating vectors. Based on the VECM, we found evidence that market power is on the demand side and retailers have a dominant position, therefore, imperfect price transmission is confirmed. In the long term, a 1 per cent increase in retail prices leads to a 0.39 per cent and 0.4 per cent increase in farm-gate and processor prices respectively. Similarly, the existing studies on the period before the end of milk quota suggest that retail prices respond asymmetrically to increases and decreases in producer prices (Weldesenbet, 2013; Lajdová *et al*., 2015). Interestingly, the findings of Bakucs *et al.* (2013) that (a) the less balanced the bargaining power of farmers and retailers, the more likely one should observe asymmetric price transmission, and (b) farm-retail price transmission asymmetry is likely to occur when retailers' turnover relative to food manufacturing turnover (per enterprise) is higher, might also explain the asymmetry in the Slovak dairy sector. However, perfect price transmission exists between farm-gate and processor prices for whole milk in the long term. Given this, the findings reveal that the recent emergence and strengthening of the producer organisations enable producers to support their bargaining position in the supply chain. The unfavourable situation after the end of milk quota, resulting in a fall in the number of milk producers, might also have contributed to the increased willingness for cooperation. There is evidence, provided by Lajdova *et al.* (2015), that opposite results held for semi-fat milk prices during the period 2003-2011, where the price adjustment from processor to producer was symmetric, but asymmetric vice versa. This may indicate that the lack of an organised contracting system before the abolition of the EU milk quota may have limited the market power of farmers relative to the dairy industry and their capacity to negotiate prices (Serra and Goodwin, 2003). The retail price movement to equilibrium is slow due to the small α-value (0.08511); meanwhile, the processor and farm-gate price restoration to the equilibrium path will take a comparatively short time. There is a two-way short-term Granger causality between processor and retail prices, farm-gate and retail prices, processor and farm-gate prices. These results are consistent with the findings of Weldesenbet (2013) and Pokrivčák and Rajčániová (2014), who conclude that the changes in producer prices cause changes in the retail prices as well as there is a causality feedback from the retail to producer prices.

We suggest the following measures in order to stabilise the dairy sector and mitigate the price asymmetry. Firstly, it is important to balance the subsidy and regulatory environment and avoid cutting off state support: the support system for the milk producers must be effective and sustainable. It is also necessary to prevent the import of milk and dairy products at dumping prices. Besides, there is also scope for improving the transparency in price formation along the supply chain; furthermore; distribution margin and the abuse of the dominant market position of retailers must be solved at the EU level.

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