Asymmetric price transmission in the Slovak liquid milk market

Tesfu WELDESENBET[1]

Agricultural Policy and Budget Section, Ministry of Agriculture and Rural Development of the Slovak Republic, Bratislava, Slovak Republic

Abstract: The divergence in liquid milk price trends has raised concern about the efficiency of the milk market price transmission in Slovakia. The paper provides empirical evidence on the price transmission among the producer, wholesale, and retail markets of liquid milk in Slovakia, using the monthly data from 1993 to 2010. The empirical analysis is based on the Granger causality and the Johansen cointegration tests and on the asymmetry tests (Houck approach and error correction model approach). The causality test results show that the changes in producer prices cause changes in the wholesale and retail prices; there is a feedback from the retail to producer prices. Moreover, the direction of causality between the wholesale and retail prices flows in both directions. The long-run elasticities of price transmission are, as expected, greater than the short-run elasticities. The cointegration results indicate that the wholesale and producer prices as well as the retail and producer prices are cointegrated, but there is no evidence of cointegration between the wholesale and retail prices. The results of an asymmetric error correction models suggest that the price transmission in the Slovakian liquid milk market is asymmetric both in the short- and long-runs.

Key words: cointegration, elasticity, error correction, milk, price asymmetry

Market price is the central mechanism by which the different levels of the market are linked. The relationship between the producer, wholesale and retail prices provides insights into the marketing channel efficiency and the degree of market competition. The milk pricing system in Slovakia has stimulated a public debate on the milk price formation and on the price transmission along the milk marketing channel. Generally speaking, there is a common perception among milk producers and their associations that dairy processors and retailers respond faster to the increases in the producer price of raw milk than to decreases. Similarly, dairy processors sometimes complain that the retailers react more rapidly to the wholesale price rises than falls. They argue that during the last eighteen years (1993–2010), the average retail price of liquid milk has increased by 118.5%, while the average producer price of raw milk has increased by only 39.6% and the average wholesale price by 60.6%. As a consequence, both consumption and production of liquid milk have reduced dramatically. In particular, the liquid milk consumption per capita declined by 38%, and the total liquid milk production by 37.5%. The growth potential of the milk sector is highly dependent, among others, on the pattern of prices. However, basic food, such as milk, shows a large variation in prices over a short period of time. For instance, between January 2010 and December 2010, the average produce price of raw milk increased sharply by 18.4% per litre, the average wholesale price of liquid milk, in turn, declined slightly by 0.5%. However, in the same period the average retail price of liquid milk increased by 10%. This situation poses two different questions. First, whether the retailers respond quickly and/or completely, when the producer price decreases? And, second, why the wholesale prices remained unchanged or even decreased slightly, when the wholesalers' input cost rises (raw milk price)? To address these questions and to better understand the milk pricing system, we study the price transmission in the liquid milk marketing channel.

The objective of this paper is to provide empirical evidence on the vertical price transmission between the Slovakian liquid milk prices at the producer, wholesale (processor) and retail levels using different econometrics analysis. In particular, to determine the existence of a long-run equilibrium relationship between a pair of prices and the speed of adjustment towards the long-run equilibrium; to estimate the

[1]The opinions expressed in this paper are those of the author and do not necessarily reflect the opinions of the Ministry of Agriculture and Rural Development of the Slovak Republic.

short-run and long-run price transmission elasticities and to investigate the direction of price causality and asymmetry in the price transmission. The empirical analysis is based on the Granger causality tests, the Johansen cointegration tests, and on two asymmetric price transmission models, i.e. the Houck's model and the asymmetric error correction model.

Price transmission might be asymmetric in its speed and magnitude, and it could differ depending on whether the price shock is positive or negative (nature), and it is being transmitted upwards or downwards (direction) along the supply chain (Vavra and Goodwin 2005). A large number of empirical studies have examined the asymmetric price transmission in milk and dairy products, either vertical or special. Kinnucan and Forker (1987) estimate the relationship between changes in the farm-level price of milk and the changes in the retail prices of four major dairy products (fluid milk, butter, cheese, and ice cream) for the period January 1971 to December 1981 in the USA using the Houck's approach. They find that the farm-retail price transmission process in the dairy sector is asymmetric. That means that the retail dairy product prices adjust more rapidly and more fully to increases in the farm price of milk than to decreases. Another study by Capps and Sherwell (2005) investigate a spatial asymmetry in the farm-retail price transmission in liquid milk products (2% milk and whole milk) for the period January 1994 to October 2002 for seven U.S. cities, applying both the Houck's approach and the error correction model. Their empirical results also suggest that the farm-retail price transmission process for milk is asymmetric. They conclude by recommending that the consideration be given to the ECM approach in addition to the Houck approach. Misra et al. (2010) also show that the cost pass through in the milk category in the USA is highly asymmetric both in the terms of the speed and magnitude. The instantaneous pass through for a cost increase is 67% as opposed to only 27% for an equivalent cost decrease. Furthermore, the evidence suggests that the asymmetry in the magnitude of the cost pass through persists in the long run. Serra and Goodwin (2002) examine the price transmission in the Spanish dairy sector using the threshold vector error correction models. They find that asymmetries are not present in the price transmission of highly perishable dairy products (pasteurized liquid milk). They reasoned that liquid milk is a product characterized by a low value added. This may tighten the relationship between the retail and farm prices and could prevent the asymmetric price adjustments. Amador et al. (2010) assess the price transmission mechanism between the producer and consumer prices of milk and dairy products in Austria using monthly data for the period January 1996 to February 2010. The results show that asymmetries play role in the milk and dairy markets in Austria. 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 the retailers. Bakucs and Fertő (2008) use the asymmetric error correction model to study the price transmission in the Hungarian milk market, and found price transmission asymmetries both in the long-run and short-run, i.e. the producer price increases are transmitted more rapidly and fully to the consumer level than the producer price decreases. As mentioned above, different authors use different methods to test the asymmetric price transmission. Frey and Manera (2007) provide reviews of developments in the econometric models of the asymmetric price transmission.

Studies on the price transmission have highlighted several causes of price asymmetries. For example, the evidence of the market power abuse or market failure is considered to be the primary potential cause of the asymmetric price transmission (von Cramon-Taubadel and Meyer 2000). Government interventions to support producer prices could also cause asymmetry (Kinnucan and Forker 1987). Ward (1982) suggests that asymmetries may be found in the price transmission of highly perishable products. If the producer prices increase, the wholesalers and retailers with perishable products would be reluctant to increase prices because they risk a lower demand and ultimately being left with the spoiled product. Another factor that affects the price transmission is the menu costs or adjustment costs (Goodwin and Holt 1999). Menu costs resulting from re-pricing or re-tagging products on the shelf and adopting a new pricing strategy might force wholesalers and retailers to be hesitant to change their prices. Some authors relate the imperfect price transmission to the existence of adjustment costs. The imperfect price transmission might exist when price changes at one end of the supply chain are not immediately reflected at the other end. The reasons are that price changes are not fully transmitted along the marketing chain; or because the increases or decreases at one end of the chain are not transmitted instantaneously, but instead distributed over time; or because the price reaction is different for the positive and negative shocks (London Economics 2004). Finally, an empirical study of the milk supply chain in Denmark (Jensen and Møller 2007) concluded that the asymmetries in the price transmission take place in the short run, whereas the price transmission is symmetric in the long run. They consider that the short-run asymmetries are the result of transaction costs, where the menu costs might be one of the most important ones in the retail sector. Long-run asymmetries might be the result of imperfections in the market structure.

A brief overview of the Slovak dairy sector

The raw milk production in the EU is regulated by quotas in order to reduce overproduction. The national milk quota of Slovakia is around 1.1 milliard kilograms. Slovakia never exceeded its national milk quotas from the beginning of the 2004/2005 quota year. From 1993 to 2010, the number of dairy cows in Slovakia decreased from 386 000 to 161 300, while the average milk yield per 1 dairy cow increased from 3042 to 5692 kg. Despite this substantial milk yield growth, the total volume production of raw milk declined from 1250 to 918 million kilograms. The majority of milk production in Slovakia is taking place in large-scale farms (co-operatives or enterprises). According to the World Bank study (2006), out of 1200 dairy farms, almost one half of them have a herd size between 100 and 500 cows. The national average per 1 dairy farm is 183 cows. However, milk production is not profitable without subsidies (Chrastinová and Burianová 2009). The dairy processing sector is, in contrast, much less concentrated. In 2003, the Slovakian top three dairy companies had the market share of less than 30%.

MATERIAL AND METHODS

In this section, two approaches to studying the price transmission process between a pair of price variables are presented. Following Capps and Sherwell (2005), first we discuss the Houck's approach. Next, the error correction model (ECM) is presented.

The Houck approach

This approach was originally introduced by Wolffram (1971) and later adapted by Houck (1977) and Ward (1982). Houck (1977) developed a test for the asymmetric price transmission based on the price segmentation technique into increasing and decreasing phases. Houck's static asymmetric model can be written as:

$$\Delta RP_t^* = RP_t - RP_0 = \beta_0 + \beta_1 \Delta WP_t^+ + \beta_2 \Delta WP_t^- + \varepsilon_t \quad (1)$$

where ΔRP_t^* is the difference between the retail price in the month t (RP_t) and the retail price in the month 0 (RP_0)[2], which is the initial month; ΔWP_t^+ is the sum of all month-to-month increases in the wholesale price (WP) from its initial value in the month 0 to the month t; ΔWP_t^- is the sum of all month-to-month decreases in the wholesale price from its initial value in the month 0 to the month t; Δ is the difference operator; t is the current time; β_0, β_1, β_2 are coefficients, and ε is the error term.

$$WP_t^+ = \begin{cases} WP_t & \text{if } WP_t > WP_{t-1} \\ 0 & otherwise \end{cases}$$

and

$$WP_t^- = \begin{cases} WP_t & \text{if } WP_t < WP_{t-1} \\ 0 & otherwise \end{cases}$$

The impact of increasing or decreasing wholesale prices on the retail price can be distributed over time. Ward (1982) extended the Houck's static asymmetric model (Equation 1) to a dynamic representation in order to include the distributed lag effect of the positive and negative cumulative price variations:

$$\Delta RP_t^* = RP_t - RP_0 = \\ = \beta_0 + \sum_{i=0}^{k} \beta_{1i} \Delta WP_{t-i}^+ + \sum_{i=0}^{l} \beta_{2i} \Delta WP_{t-i}^- \quad (2)$$

The β_1 and β_2 coefficients in Equation (2) measure the impact of rising and falling wholesale prices on retail prices, respectively. k is the lag length for increasing prices; l is the lag length for deceasing prices. The null hypothesis of the long-run symmetry (H_0) and the alternative hypothesis of the long-run asymmetry (H_A) tests in Equation (2) are as follow:

$$H_0: \sum_{i=0}^{k} \beta_{1i} = \sum_{i=0}^{l} \beta_{2i} \qquad H_A: \sum_{i=0}^{k} \beta_{1i} \neq \sum_{i=0}^{l} \beta_{2i} \quad (3)$$

The meaning of the null hypothesis is that the cumulative effect of increases in the wholesale price should be equivalent to the cumulative effect of decreases in the wholesale price. A rejection of H_0 is the evidence of asymmetry that means that the retailers' responses to the wholesale price increases differ from the responses to the wholesale price decreases. A fail to reject H_0 is the evidence of symmetry in the price transmission. This hypothesis can be evaluated us-

[2]Kinnucan and Forker (1987) demonstrate examples of the price date transformation in order to be studied using the Houck's approach.

ing the F-test. The null hypothesis of the short-run symmetry is rejected when the individual β_{1i} and β_{2i} coefficients are unequal.

Cramon-Taubadel (1998) point out that the Houck's approach ignores the important time-series properties of the data. Estimating Equation (2) without knowing the time-series properties of price data can lead to a spurious regression problem and the results will be misleading. Since prices are usually non-stationary and cointegrated, it is desirable to conduct the preliminary data analysis and to utilize the long-run information that might provide price data in levels.

Error correction model

The error correction model (ECM) proposed by Engle and Granger (1987) suggest a two-step procedure. First, a static cointegration regression is estimated and tests for cointegration are conducted. If the producer prices (PP) and retail prices (RP) are cointegrated, assuming that the producer prices lead retail prices, then their long-term equilibrium relationship can be expressed as follows:

$$RP_t = \beta_0 + \beta_1 PP_t + \varepsilon_t \tag{4}$$

Second, if the residuals from the cointegration regression (4) are stationary, i.e. the cointegration, the lagged residuals are imposed as the error correction term (ECT_{t-1}) in the ECM:

$$\Delta RP_t = RP_t - RP_{t-1} = \beta_0 + \sum_{i=0}^{k} \beta_{1i} \Delta PP_{t-i} + \sum_{i=1}^{l} \beta_{2i} \Delta RP_{t-i} + \beta_3 ECT_{t-1} + \varepsilon_t \tag{5}$$

Where ECT_{t-1} is the one period lagged deviation from the long-run equilibrium Equation (4) between the retail and producer prices $ECT_{t-1} = \varepsilon_{t-1} = RP_{t-1} - \beta_0 - \beta_1 PP_{t-1}$). Equation (5) is the specification of the ECM with a symmetric adjustment of deviations from the long-term equilibrium. β_3 measures the adjustments to deviations from the long-term equilibrium, while the short-term dynamics is measured by β_{1i} and β_{2i}, $i = 1, ..., k$, and $i = 1, ..., l$, respectively. Von Cramon-Taubadel (1998) proposed a modification to Equation (5) that involves the segmentation of the error correction term (ECT) into its positive (ECT^+) and negative (ECT^-) components. Furthermore, Von Cramon-Taubadel and Loy (1999) made an additional modification to Equation (5) by segmenting the producer price into its positive (PP^+) and negative (PP^-) segments. The modified version of Equation (5) can be written as:

$$\Delta RP_t = \beta_0 + \sum_{i=0}^{k1} \beta_{1i}^{+} \Delta PP_{t-i}^{+} + \sum_{i=0}^{k2} \beta_{2i}^{-} \Delta PP_{t-i}^{-} + \sum_{i=1}^{l} \beta_{3i} \Delta RP_{t-i} + \beta_4^{+} ECT_{t-1}^{+} + \beta_4^{-} ECT_{t-1}^{-} + \varepsilon_t \tag{6}$$

$$ECT_t^{+} = \begin{cases} ECT_t & \text{if } ECT_t > 0 \\ 0 & otherwise \end{cases}$$

and

$$ECT_t^{-} = \begin{cases} ECT_t & \text{if } ECT_t < 0 \\ 0 & otherwise \end{cases}$$

The null hypothesis of the symmetry test in Equation (6) becomes as follows:

$$H_0: \sum_{i=0}^{k1} \beta_{1i}^{+} = \sum_{i=0}^{k2} \beta_{2i}^{-} \quad \text{and} \quad \beta_4^{+} = \beta_4^{-} \tag{7}$$

The symmetry or asymmetry of the price transmission in Equation (7) can be evaluated using a joint F-test.

DATA

The average monthly data for the producer, wholesale and retail prices of liquid milk are used for estimation. The time-series data cover the period from January 1993 to December 2010, giving a total of 216 observations. Data for average producer prices of raw milk and wholesale prices of liquid milk have been obtained from the Ministry of Agriculture and Rural Development of the Slovak Republic. The Statistical Office of the Slovak Republic provided data for average retail prices of liquid milk. Pasteurized semi-fat milk (1.5% fat) bagged in plastic is selected for analysis. The advantage of using this type of milk for analysis is its relative homogeneity. This product compared to other dairy products undergoes only light processing (pasteurization, homogenization and fat content adjustment) between the farm and the retail market. Thus, a close relationship among three prices is expected. Since Slovakia joined the monetary union on 1 January 2009, prices are expressed in Euros after 2009. Note that prices of year 2009 and 2010 are converted to the previous currency (Slovak Crown, SKK) using the official conversion rate (30.1260 SKK/1 Euro). As a result, all milk prices series in this paper are expressed in Slovak Crowns per litre

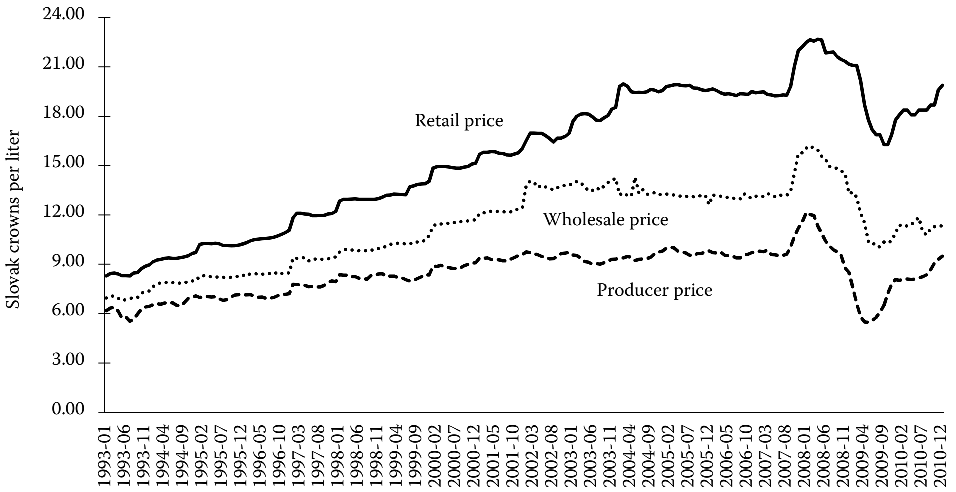

Figure 1. Development of producer, wholesale and retail prices of liquid milk in Slovakia, January 1993 to December 2010

(in SKK per litre). The econometric software EViews is used to run the models.

RESULTS AND DISCUSSION

The liquid milk market prices have shown a considerable variability and an upward trend over the past eighteen years in Slovakia (Figure 1). The Figure indicates that although there is an upward trend in all prices, the pattern of price movements as well as their extent is differing. While the upward common trend might suggests the existence of a cointegration relationship, the extent of price movements might show the presence of asymmetric price transmission. The Figure 1 also indicates that the highest average producer price of raw milk in history was in January 2008 (12.16 SKK/litre); average wholesale price in February 2008 (16.20 SKK/litre); and average retail price in April 2008 (22.68 SKK/litre) due to milk supply imbalance in the world market.

Descriptive statistics of liquid milk prices are shown in Table 1. Analysis of price dispersion in liquid milk prices is also described by using the coefficient of variation. A higher value of the coefficient of variation indicates a larger dispersion of monthly prices and vice versa. In other words, higher dispersion implies the more uncertain the price was due to a higher degree of variation around the mean. This analysis is conducted on the entire sample period (1993–2010) as well as on sample sub-periods, before Slovakia joining the EU (1993–April 2004) and after Slovakia joining the EU (from 1st of May 2004 onwards) to identify price variations across time (Table 1). Over the entire sample period, average retail price was more volatile (coefficient of variation 26.4%) than wholesale price, and wholesale price was more volatile than producer price. In contrast, after Slovakia joining the EU, producer milk price was more volatile (coefficient of variation 15.7%) than the wholesale price and wholesale price (11.7%) was more volatile than the retail price (7.5%).

Unit root tests

In order to determine the level of integration of the individual price series the Dickey-Fuller Generalised

Table 1. Descriptive statistics of liquid milk prices in Slovakia, 1993–2010 (SKK/litre)

Price	Obs.	Mean	Std. dev.	Min.	Max.	CV (%)		
						1993–2010	before joining the EU	after joining the EU
Producer	216	8.465	1.375	5.480	12.164	16.2	14.3	15.7
Wholesale	216	11.297	2.395	6.790	16.200	21.2	22.0	11.7
Retail	216	15.567	4.110	8.290	22.680	26.4	24.2	7.5

CV is coefficient of variation = (standard deviation/mean) × 100

Source: own calculations

Table 2. Unit root tests results

	Level			First differences		
	none	constant	constant and trend	none	constant	constant and trend
Producer price						
PP	0.434	–2.342	–2.544	–7.045***	–7.047***	–7.056***
DF-GLS	–	–1.153 (1)	–3.344**(2)	–	–4.204**(1)	–6.400** (0)
Wholesale price						
PP	0.897	–2.080	–1.082	–12.967***	–13.018***	–13.163***
DF-GLS	–	0.065 (0)	–0.438 (0)	–	–6.676*** (1)	–12.828*** (0)
Retail price						
PP	2.206	–1.979	–1.439	–8.354***	–8.657***	–8.731***
DF-GLS	–	0.823 (1)	–1.370 (1)	–	–6.739***(0)	–8.368*** (0)

The critical values for *PP* tests with constant, constant and trend and none at the 5% significance are –2.87, –3.43 and –1.94. The critical values at the 1% significance are –3.46, –4.00 and –2.57. The critical values for DF-GLS tests with constant, constant and trend at the 5% significance are –1.94, and –2.93. The critical values at the 1% significance are –2.57 and –3.46. The tested hypothesis (PP and DF-GLS tests) is H_0: non-stationary and H_A: stationary
*** and ** indicate significance at the 1% and 5% levels. Lag length selection is automatic based on Schwarz Information Criteria. Lag length in parentheses

Source: own calculations

Least Squares (DF-GLS) proposed by Elliott et al. (1996) and Phillips Perron (PP) unit root tests are used. To cross-check the results of PP and DF-GLS tests we applied also KPSS (Kwiatkowski, Phillips, Schmidt and Shin 1992) test. Non-stationarity is the null hypothesis in the DF-GLS and PP tests, while stationarity is the null hypothesis in the KPSS test. For the analysis we take natural logarithms of all price series. Table 2 reports the results of the unit root tests in level and first difference. For each price variable, the results fail to reject the null hypothesis of non-stationarity in the level at the 5% significance level. The only exception is milk producer prices that the DF-GLS test indicated stationary in level. However, the PP and KPSS (Table 3) tests to detect a unit root in the level of producer prices. When we take the first differences of each price series, both DF-GLS and PP tests rejected the null hypothesis of non-stationarity. Therefore, we conclude that all three milk price series are integrated of order one, i.e. I (1), meaning that they can be stationary by first differencing. The next step is to examine the cointegration relationships between these variables.

Pairwise Granger causality tests

Before proceeding to estimating the asymmetric price transmission equations, Granger causality tests are conducted to determine the causal linkages

Table 3. Stationarity tests results

	Level		First differences	
	constant	constant and trend	constant	constant and trend
KPSS – Producer price	0.926	0.290	0.059***	0.030***
KPSS – Wholesale price	1.428	0.372	0.341***	0.043***
KPSS– Retail price	1.741	0.379	0.279***	0.031***

The tested hypothesis (KPPS test) is H_0: stationary and H_A: non-stationary. The critical values for KPSS tests with constant, constant and trend at the 5% significance are 0.463 and 0.146. The critical values at the 1% significance are 0.739 and 0.216
*** indicates significance at the 1% level

Source: own calculations

Table 4. Pairwise Granger causality tests

Null hypothesis:	Obs.	*F*-statistic	Probability	Lags
ΔPP does not Granger cause ΔRP	213	6.692	0.000***	2
ΔRP does not Granger cause ΔPP	213	22.043	0.000***	2
ΔWP does not Granger cause ΔPP	210	1.116	0.353	5
ΔPP does not Granger cause ΔWP	210	14.012	0.000***	5
ΔWP does not Granger cause ΔRP	214	7.562	0.007***	1
ΔRP does not Granger cause ΔWP	214	27.388	0.000***	1

*** indicates rejection of null hypothesis at 1% significance level

Source: own calculations

between two price variables. A necessary condition for Granger causality tests is that each of the price series is stationary. The unit root tests described above found that all price series in level are non-stationary, therefore the analysis of Granger causality tests are conducted on first differentials. The basic idea of Granger (1969) causality theory is to test the null hypothesis that changes in one variable are not able to predict the other. Granger causality tests allow us to make some inferences about the direction of information flows between two price series. Since the Granger causality test is very sensitive to the number of lags included in the regression, both the Akaike Information Criterion (AIC) and Schwarz Information Criterion (SIC) are used in order to find an appropriate number of lags. The causality tests results are presented in Table 4. The results show that the null hypothesis that change in the producer prices does not Granger cause change in the wholesale and retail prices is rejected at 1% significance level. There is a feedback from retail to producer prices. However, we fail to reject the null hypothesis that change in the wholesale prices does not Granger cause change in the producer prices at any significance levels. This implies that changes in the producer prices of raw milk clearly lead changes in the wholesale prices and not vice versa. The results indicate that dairy farmers in Slovakia are not price takers but can influence market prices. In the UK liquid milk market, for instance, the causation of price setting is from wholesaler to farmer and not vice versa (Shabbar and Grigoryev 2011). Furthermore, the causality direction between wholesale and retail prices flows in both directions. This means changes in the price of liquid milk in one market will affect the other. These two prices are interdependent and there is no clear price leader and follower between them. This result is in consistent with findings of Shabbar and Grigoryev (2011), who also found bi-directional causality between wholesale and retail prices of liquid milk in the UK.

Cointegration

The results, shown in Tables 2 and 3, reveal that all price series are integrated of order one. The next step then involves testing for the existence of long-run equilibrium relationships between a pair of price series-using Johansen cointegration tests. The concept of cointegration (Engle and Granger 1987) states that individual economic variables may be non-stationary and wonder through time, but a linear combination of them may, over time, converge to a stationary process. Such a process, if present, may reflect the long run equilibrium relationship. In other words, cointegration implies that the two integrated prices never drift apart from each other in the long-term, i.e. they maintain equilibrium relationship. However, the cointegrated prices can deviate temporarily but market forces bring the prices to their cointegrated equilibrium. The Johansen procedure results are sensitive to the data lag structure. Therefore, we estimate a vector autoregression (VAR) using a pair of price level series, and then based on different statistical tests (Akaike, Schwarz and Hannan-Quinn Information Criteria) criteria selected the optimum lag lengths. The results of the cointegration analysis are presented in Table 5. As Table 5 shows, there is no evidence of cointegration between wholesale and retail price series of liquid milk. Both trace and maximal-eigenvalue statistics fail to reject the null hypothesis of no cointegration at the 5% significance level. This indicates that wholesale and retail milk prices are separated and drift apart significantly in the long run. The absence of long run or equilibrium relationship between wholesale and retail prices could be interpreted as some evidence against

Table 5. Johansen cointegration tests

Liquid milk	Lag length	Hypothesized of cointegrated equation	Trace statistic	P-value	Maximal eigenvalue statistic	P-value
PP-WP	2	none	35.893	0.000***	32.128	0.000***
		at most 1	3.765	0.052	3.765	0.052
PP-RP	3	none	22.651	0.004***	19.042	0.008***
		at most 1	3.609	0.058	3.609	0.058
WP-RP	2	none	5.850	0.713	3.997	0.860
		at most 1	1.854	0.173	1.854	0.173

The intercept (no trend) option is selected. *** indicates a significant level at 0.01

Source: own calculations

market competitive. Different results were found in the UK and Greek's liquid milk markets where wholesale and retail prices are cointegrated (Reziti 2005; Shabbar and Grigoryev 2011). Furthermore, our results indicate that wholesale and retail prices are cointegrated with producer prices of raw milk, i.e. there are co-movements in these price series over a long-term. The null hypothesis of no cointegration is even rejected strongly at 1% significance level. Similar results were found for the U.S. (Stewart and Blayney 2010) and Poland (Bakucs et al. 2012) where retail and farm prices for fluid milk are cointegrated. On the other hand, Reziti (2005) has found no evidence for long-run relation between farmers and wholesalers in Greek's milk market.

Houck approach results

In the previous section we failed to find evidence of cointegration between retail and wholesale prices of liquid milk. Hence, Houck approach is estimated using Equation (2). From a statistical point of view, there are a sufficient number of observations in wholesale prices decreases (44%) and increases (56%). Estimating a distributed lag model such as Equation (2) might cause a multicollinearity problem since the lagged wholesale prices are likely correlated with each other. To solve this problem we applied the polynomial distributed lag (PDL) model, also known as the Almon lag model (Gujarati 2003). The Almon model is used to estimate the short-run and long-run impacts of wholesale price changes on retail prices. Lag structures are assumed to lie on a second order polynomial with endpoint constraints. The lag lengths are determined by the Schwarz Information Criterion (SIC). We found that the lag length of two ($k1 = k2 = 2$) gives the lowest SIC statistic. This reveals that response to wholesale prices could likely occur within two months. Since significant serial correlation was evident in this equation estimation, we present generalized least squares estimates.

The results of the coefficient estimates based on the Houck approach are reported in Table 6. The estimated coefficients have the expected signs and are highly significant. The exception is the coefficient of ΔWP_t^-, which is statistically insignificant.

Table 6. Empirical results of the Houck approach

Variable	Dependent variable (ΔRP_t^*)	
	estimated coefficient	t-statistic
Constant	1.703	1.963
ΔWP_t^-	0.061	0.753
ΔWP_{t-1}^-	0.281	5.685***
ΔWP_{t-2}^-	0.261	4.552***
$\Sigma\Delta WP_t^-$	0.602	5.443***
ΔWP_t^+	0.574	8.054***
ΔWP_{t-1}^+	0.302	6.540***
ΔWP_{t-2}^+	0.111	2.145**
$\Sigma\Delta WP_t^+$	0.988	9.450***
AR(1)	1.243	17.289***
AR(2)	0.280	–3.945***
R^2	0.997	
DW	1.943	
AIC	–0.199	
SIC	–0.088	
P-value (J-B)	0.000	
P-value (B-G)	0.465	
P-value (ARCH)	0.583	

Lag length is determined by SIC criterion. *** and ** indicate significant at 1% and 5%, respectively

Source: own calculations

Table 7. Price transmission elasticities between retail and wholesale prices of liquid milk in Slovakia

Price increases		Price decreases	
Short-run	long-run	short-run	long-run
0.417	0.717	0.044	0.437

Source: own calculations

This indicates that the immediate effect of wholesale price decreases on retail prices is null. However, there are significant effects in the following months. The coefficient of determination is 0.997, implying that almost all (99.7%) retail price fluctuations are explained by wholesale price changes. The results also show that the immediate response of the retail price to wholesale price change is greater for increases than decreases. For example, a 10 hallers (cents) per litre increase in the wholesale price results in a 5.7 hallers increase in the retail price, while a 10 hallers decrease in wholesale prices results in the retail price decrease only 0.6 hallers. Both price shocks are transmitted only incompletely ($\Delta WP_t^- < \Delta WP_t^+ < 1$). In percentage terms, the largest portion 58% from the complete pass through of the wholesale price increases occurs in the immediate time period, compared to 10% from an equivalent wholesale price decreases. In addition, according to Wald test, the null hypothesis that the cumulative effect of increases in wholesale price (ΔWP_t^+) is equivalent to the cumulative effect of decreases (ΔWP_t^-) in wholesale price (Equation 3) is rejected strongly. In sum, the results over the period (1993–2010) suggest asymmetric price transmission exists between wholesale and retail prices of liquid milk in Slovakia. This means retail prices respond more strongly to wholesale price increases than to wholesale price decreases. In other words, price transmission between retail and wholesale prices of liquid milk is positive. This result is not in line with Ward's (1982) result that retail prices respond more to a wholesale price decrease than a wholesale price increase. He reasoned that increasing prices might reduce retail sales and increase the spoilage of highly perishable products.

Furthermore, Table 6 provides some diagnostic tests for normality (Jarque-Bera test), heteroscedasticity (ARCH test), and autocorrelation (Breusch-Godfrey test). The results illustrate some problems with normality in the price transmission equation, but overall the model specifications seem not to suffer from problems of autocorrelation and heteroskedasticity.

Elasticities of price transmission, calculated at the sample means of the retail and wholesale prices, are presented in Table 7. The elasticity of price transmission is defined as the percentage changes in retail price due to a one percent change in wholesale price. We calculated both the short-run and long-run elasticities of price transmission for increasing and decreasing wholesale prices. The results confirm that the long-run elasticities of price transmission, as expected, are greater than the short-run elasticities of price transmission. This result holds for both increases and decreases in prices at the wholesale level. For instance, one percent increase in wholesale prices in a short-run leads 0.417% increase in retail prices, while one percent decrease in wholesale prices in short-run causes only 0.044% reduction in retail prices. Of course, elasticities tend to increase over time as retailers have more options to search for better supply deals.

Error correction model results

The cointegration of wholesale and retail prices with producer prices enables us to estimate symmetric and asymmetric error correction models (ECM) between these prices. Von Cramon-Taubadel and Loy (1999) have demonstrated that application of the asymmetric ECM is more appropriate than the use of the Houck approach if the price data are cointegrated. Tables 8 and 9 summarize the results of the long-run equilibrium regressions and of the symmetric and asymmetric error correction models. The results of the long-run equilibrium regressions indicate that all the estimated coefficients have the correct sign are statistically highly significant. A 1% rise in the producer prices of milk results in increases of retail milk prices by 2.352%, while the rise of the same magnitude in the producer prices would increase wholesale prices by 1.551%. Thus, producer price increases have a greater impact on wholesale and retail prices. Similarly, price decreases in producer prices produce a proportionally greater decrease in both prices.

Next we will analysis an error correction terms in the ECMs, from which the speed of adjustments towards the long-run values are estimated. The number of lags in Equation (5) is determined by the Schwarz Information Criteria (SIC). In case of retail-to-producer prices we found that three lagged values ($k1 = k2 = 3$, i.e. three months) yield the lowest SIC statistic for the estimated error correction models. In case of wholesale-to-producer prices we found two lagged values ($k1 = k2 = 2$, i.e. two months). Table 8 presents the results of symmetric error correction models. With this symmetry equation the price increases have the same absolute impact as the price decreases on the dependant variable. The results indicate that the R-squared statistics are low (0.45 and 0.32). The reason for this is that the dependent variables in ECMs are estimated in differenced form, i.e. changes in wholesale and changes in retail prices. Furthermore, the coefficients of the error correction terms have the expected negative sign and are both significant at 1%. The coefficient of β_3 is expected to be negative because when prices are above (below) equilibrium, prices are expected to decrease (increase). According to our results the deviations from the equilibrium levels of wholesale and retail prices during the current month are corrected at about 8% and 2% in the next month, respectively. It

Table 8. Symmetric error correction models results

Variable	Dependent variable ΔRP		Dependent variable ΔWP	
	estimated coefficient	t-statistic	estimated coefficient	t-statistic
Constant	0.033	2.218**	0.015	0.889
ΔPP_t	0.232	3.144***	0.271	2.794**
ΔPP_{t-1}	0.205	6.641***	0.282	4.583***
ΔPP_{t-2}	0.157	2.955***	0.192	2.406**
ΔPP_{t-3}	0.088	1.916	–	–
$\Sigma\Delta PP_t$	0.682	6.641***	0.746	6.466***
ΔRP_{t-1}	0.302	4.393***	–	–
ΔRP_{t-2}	–0.090	–1.364	–	–
ΔWP_{t-1}	–	–	–0.151	–2.288**
ΔWP_{t-2}	–	–	–0.068	–1.070
ECT_{t-1}	–0.018	–2.813***	–0.078	–4.275***
R^2	0.447		0.321	
DW	2.014		1.992	
AIC	–0.227		0.061	
SIC	–0.132		0.156	
P-value (J-B)	0.000		0.000	
P-value (B-G)	0.473		0.305	
P-value (ARCH)	0.781		0.066	
Cointegration regressions				
	Dependent variable RP		Dependent variable WP	
	estimated coefficient	t-statistic	estimated coefficient	t-statistic
Constant	–4.344	–4.021***	–1.829	–3.933***
	2.352	18.671***	1.551	28.599***

Source: own calculations

Table 9. Asymmetric error correction models results

Variable	Dependent variable ΔRP		Dependent variable ΔWP	
	estimated coefficient	t-statistic	estimated coefficient	t-statistic
Constant	0.076	2.344	0.027	0.821
ΔPP_t^-	–0.316	–2.755**	–0.014	–0.090
ΔPP_{t-1}^-	0.055	1.216	0.119	1.414
ΔPP_{t-2}^-	0.231	3.014**	0.124	1.093
ΔPP_{t-3}^-	0.213	3.153**	–	–
$\Sigma\Delta PP_t^-$	0.183	1.216	0.230	1.425
ΔPP_t^+	0.141	1.260	0.072	0.498
ΔPP_{t-1}^+	0.324	6.369***	0.522	6.394***
ΔPP_{t-2}^+	0.362	4.942***	0.497	4.798***
ΔPP_{t-3}^+	0.254	4.045***	–	–
$\Sigma\Delta PP_t^+$	1.080	6.369***	0.091	6.407***
ΔRP_{t-1}	0.230	3.383**	–	–
ΔRP_{t-2}	–0.136	–2.148**	–	–
ΔWP_{t-1}	–	–	–0.159	–2.557**
ΔWP_{t-2}	–	–	–0.066	–1.082
ECT_{t-1}^+	–0.061	–4.998***	–0.161	–5.928***
ECT_{t-1}^-	0.036	1.863	0.012	0.240
R^2	0.509		0.374	
DW	1.985		2.000	
AIC	–0.316		0.013	
SIC	–0.173		0.156	
P-value (J-B)	0.000		0.000	
P-value (B-G)	0.828		0.799	
P-value (ARCH)	0.926		0.207	
Hypothesis	Long-run		Short-run	
Symmetry	$\beta_4^+ = \beta_4^-$		$\sum_{i=0}^{k1}\beta_{1i}^+ = \sum_{i=0}^{k2}\beta_{2i}^-$	
Asymmetry	$\beta_4^+ \neq \beta_4^-$		$\sum_{i=0}^{k1}\beta_{1i}^+ \neq \sum_{i=0}^{k2}\beta_{2i}^-$	
Long-run & short-run symmetry tests results				
Long-run	F (1,202) = 12.541 P-value (0.000)***		F (1,203) = 7.095 P-value (0.008)***	
Short-run	F (1,202) = 26.289 P-value (0.000)***		F (1,203) = 5.127 P-value (0.025)**	
Cointegration regressions				
	Dependent variable RP		Dependent variable WP	
	estimated coefficient	t-statistic	estimated coefficient	t-statistic
Constant	–4.344	–4.021***	–1.829	–3.933***
PP	2.352	18.671***	1.551	28.599***

***1% significance level and**5% significance level

Source: own calculations

seems that prices appear to adjust relatively faster in wholesale than retail level.

Table 9 provides the results of asymmetric error correction models. Coefficients $\beta_4^+$and β_4^- asymmetric adjustment speeds, which measures long-run asymmetry, while the coefficients (β_{1i}^+ and β_{2i}^-), $i = 1, ..., k$, measures short-run asymmetry. The estimated coefficients indicate that positive error correction terms (ECT_{t-1}^+) are significant at 1% level, while the negative error correction terms (ECT_{t-1}^-) are insignificant at 5% level. This implies that the retail and wholesale prices of liquid milk are above their equilibrium prices with respect to the produce prices of raw milk so we would expect the retail and wholesale prices to adjust downward in the following month. In particular, positive deviations of the average retail and wholesale prices from the long-term equilibrium are reduced by 6.1% and by 16.1% per month, respectively. The null hypothesis (long-run symmetric price transmission) that retail and wholesale prices respond symmetrically to increases and decreases in producer prices in the long-run Equation (7) is rejected. These asymmetric results indicate that retailers and wholesalers react more quickly to producer price increases than to declines. Likewise, short-run estimates show that after two or three months the effects of producer price increases are larger than those of producer price decreases. The *F*-values of tests on short-run symmetry are high, meaning that price transmission is asymmetric also in the short-run (Table 9). Interestingly, these results are in line with milk producers' perception of the actual effects of producer price variations on retail and wholesale price changes. Bakucs et al. (2012) have found similar short- and long-term price asymmetries in Polish milk market. Finally, some diagnostic tests are used to check for residuals normality, serial correlation and heteroskedasticity in both error correction models (Tables 8 and 9). The Jarque-Bera (J-B) tests for normality of the residuals in both equations suggest that the residuals are not normal distributed. However, the Breusch-Godfrey (B-G) tests fail to reject the null hypothesis of no autocorrelation in both models. Finally, the ARCH tests indicate the absence of heteroskedasticity in the models.

CONCLUSION

This paper empirically investigated the vertical price transmission among producer, wholesale and retail prices of liquid milk in Slovakia. The data used in this analysis are the average monthly price series from 1993 to 2010. Unit root tests suggest that all the milk price series are stationary in first differences, and hence are first-order integrated. The Granger causality tests indicate that producer price changes are a significant predictor of changes in wholesale and retail prices of liquid milk in Slovakia. We found that there is a causality feedback from retail prices, meaning that changes in the retail prices cause changes in the raw milk prices, too. However, producer prices do not adjust to wholesale market shocks. Furthermore, the causality direction between wholesale and retail prices flows in both directions. The results of cointegration analysis reveal that the wholesale and retail prices of liquid milk move together with producer price in the long-run, that is, they are cointegrated. However, there is no evidence of cointegration between wholesale and retail prices. Therefore, to analyse the relationship between retail and wholesale prices we applied the Houck's approach. The results indicate that the magnitude of the wholesale price increases transmission to the retail price exceeds the magnitude of the wholesale price decreases both in short-run and long-run, i.e. asymmetric price transmission. Moreover, price transmission elasticities between retail and wholesale prices are found to be inelastic, but price transmission elasticities of rising wholesale prices are larger than falling wholesale prices. For cointegrated series, the results of error correction models suggest that both retail and wholesale prices of liquid milk adjust more quickly to increases in the producer prices, but adjust more slowly to decreases in producer prices, i.e. asymmetric price transmission. The efficiency of the sector is reducing if processors and retailers delay and not completely transmit producer price decreases to the wholesale and consumer level. In such a case, as we found, implies a welfare loss for consumers because the consumers do not benefit from producer price declines; they only absorb price increases in the milk sector. In general, the results confirm the current market sentiments of both producers and wholesalers.

REFERENCES

Amador O.F, Baumgartner J., Cuaresma J.C (2010): Milking the Prices: The Role of Asymmetries in the Price Transmission Mechanism for Milk Products in Austria. Working Papers 378. WIFO, Wien, Available at http://www.wifo.ac.at/wwa/servlet/wwa.upload.DownloadServlet/bdoc/WP_2010_378$.PDF (accessed April 2012).

Bakucs L.Z., Fertö I. (2008): Price Transmission on the Hungarian Milk Market. In: 12th EAAE Congress 'People, Food and Environments: Global Trends and European Strategies'. Gent, 26–29 August 2008. Available at http://ageconsearch.umn.edu/bitstream/44175/2/330a.pdf (accessed May 2012).

Bakucs Z., Falkowski J., Fertö I. (2012): Does Farm and Processing Industry Structure Matter for Price Transmission? Some Evidence from Transition Countries: A Comparison of Dairy Sectors in Hungary and Poland. Discussion Papers MT-DP -2012/12. IEHAS, Budapest. Available at http://econ.core.hu/file/download/mtdp/MTDP1212.pdf (accessed January 2013).

Capps O. Jr., Sherwell P. (2005): Spatial Asymmetry in Farm-Retail Price Transmission Associated with Fluid Milk Products. In: American Agricultural Economics Association Annual Meeting, Providence, Rhode Island. Available at http://future.aae.wisc.edu/publications/asymmetryAAEA05_capps_paper_final_version.pdf (accessed February 2012).

Chrastinová Z., Burianová V. (2009): Economic development in Slovak Agriculture. Agricultural Economics – Czech, *55*: 67–76.

Commission of the European Communities (2009): Analysis of price transmission along the food supply chain in the EU: Commission Staff Working Document. Available at http://ec.europa.eu/economy_finance/publications/publication16067_en.pdf (accessed May 2012).

Elliott G., Rothenberg T.J.,Stock J.H. (1996): Efficient tests for an autoregressive unit root. Econometrica, *64*: 813–836.

Engle R.F., Granger C.W.J. (1987): Co-integration and error correction: Representation, estimation and testing. Econometrica, *55*: 251–276.

Frey G., Manera M. (2007): Econometric models of asymmetric price transmission. Journal of Economics Surveys, *21*: 349–415.

Goodwin B.K., Holt M.T. (1999): Price transmission and asymmetric adjustment in the US beef sector. American Journal of Agricultural Economics, *81*: 630–637.

Granger C. (1969): Investigating causal relations by econometric models and cross spectral methods. Econometrica, *37*: 424–438.

Gujarati D.N. (2003): Basic Econometrics. 4th ed. McGraw-Hill, New York; ISBN 0-07-112342-3.

Houck P.J. (1977): An approach to specifying and estimating non reversible functions. American Journal of Agricultural Economics, *59*: 570–572.

Jensen J.D., Møller A.S. (2007): Vertical Price Transmission in the Danish Food Marketing Chain. Institute of Food and Resource Economics. Working Papers. Institute of Food and Resource Economics, Copenhagen. Available at http://www.foi.life.ku.dk/publikationer/~/media/migration%20folder/upload/foi/docs/publikationer/working%20papers/2007/wp%2008_2007.pdf.ashx (accessed Jun 2012).

Johansen S. (1988): Statistical analysis of cointegrating vectors. Journal of Economic Dynamics and Control, *12*: 231–254.

Kinnucan H.W., Forker O.D. (1987): Asymmetry in farm-retail price transmission for major dairy products. American Journal of Agricultural Economics, *69*: 285–292.

Kwiatkowski D., Phillips P.C.B., Schmidt P., Shin Y. (1992): Testing the null hypothesis of stationarity against the alternative of a unit root. Journal of Econometrics, *54*: 159–178.

London Economics (2004): Investigation of the determinants of farm-retail price spreads. Final report to DEFRA. Available at http://archive.defra.gov.uk/evidence/economics/foodfarm/reports/pricespreads/wholerep.pdf (accessed May 2012).

Misra K. Khan R., Singh V. (2010): Asymmetries and Dynamics of Cost pass through in the US Milk Market. Available at http://faculty.london.edu/kmisra/assets/documents/Cost_Passthrough.pdf (accessed April 2012).

Reziti I.N (2005): An Investigation into the Relationship between Producer, Wholesale and Retail Prices of Greek Agricultural Products. In: 79th AES Annual Conference, University of Nottingham, Nottingham, 4–6 April 2005.

Serra T., Goodwin B.K. (2002): Price Transmission and Asymmetric Adjustment in the Spanish Dairy Sector. 2002 AAEA-WAEA Annual Meeting. Available at http://sard.ruc.edu.cn/zengyinchu/files/Kecheng/Agricultural%20Market%20and%20Distribution/Research%20Papers/PRICE%20TRANSMISSION%20AND%20ASYMMETRIC%20ADJUSTMENT%20IN%20THE%20SPANISH%20DAIRY%20SECTOR.pdf (accessed May 2012).

Shabbar J., Grigoryev R. (2011): Asymmetric Price Transmission within the UK Dairy Supply Chain. University of Portsmouth, Centre for Economic Analysis and Policy. Available at http://www.port.ac.uk/departments/academic/economics/research/researchcentres/ceap/recentresearchoutput/downloads/filetodownload,137820,en.pdf (accessed May 2012).

Stewart H., Blayney D. (2010): Retail dairy prices fluctuate with the farm value of milk. In: Agricultural & Applied Economics Association 2010 AAEA, CAES, & WAEA Joint Annual Meeting, Denver, Colorado, July 25–27, 2010. Available at http://ageconsearch.umn.edu/bitstream/60947/2/AAEA%20Denver%20Poster%2310968.pdf (accessed April 2012).

Vavra P., Goodwin, B.K (2005): Analysis of Price Transmission along the Food Chain. Working Papers No. 3. OECD Food, Agriculture and Fisheries, Paris.

Von Cramon-Taubadel S. (1998): Estimating asymmetric price transmission with the error correction representation: An application to the German pork market. European Review of Agricultural Economics, *25*: 1–18.

von Cramon-Taubadel S., Loy J.P. (1999): The Identification of Asymmetric Price Transmission Processes with Integrated Time Series. Jahrbucher for Nationalokonomie und Statistik, *218*: 85–106.

Von Cramon-Taubadel S., Meyer J. (2000): Asymmetric Price Transmission: Fact or Artifact? Working Paper. Department of Agricultural Economics, Göttingen. Available at http://www.jochenmeyer.de/pdf/cramon-meyer.pdf (accessed May 2012).

Ward RW. (1982): Asymmetry in retail, wholesale, and shipping point pricing for fresh vegetables. American Journal of Agricultural Economics, *62*: 205–212.

Wolffram R. (1971): Positivistic measures of aggregate supply elasticities: Some new approaches – some critical notes. American Journal of Agricultural Economics, *53*: 356–359.

World Bank (2006): Market Linkages in the Slovak Agri-Food Sector. Working Paper No. 43; ISBN 80-89244-02-5.

Received: 28th December 2012
Accepted: 27th August 2013

Contact address:

Tesfu Weldesenbet, Ministry of Agriculture and Rural Development of the Slovak Republic, Dobrovičova 12, 812 66 Bratislava, Slovak Republic
e-mail: Tesfu.Weldesenbet@land.gov.sk