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Vertical price transmission along the diary supply chain in Russia

Many studies have analysed vertical price transmission using time-series econometric methods but vertical price transmission in the milk market in Russia has not been investigated. This paper studies vertical price transmission along the whole milk supply chain in the Russian market using the autoregressive distributed lags model. Monthly farm-gate and retail prices in Voronezh *Oblast*, a historically large agrarian region located to the south of Moscow, covering the period from 2002 to 2014 were used in the analysis. When estimating the vertical price transmission in the dairy market, seasonality should be taken into account. Using a cointegration technique, no empirical evidence is found for cointegration between farm-gate and retail prices. There is unidirectional Granger causality from retail to farm prices and not vice versa. The results support the assumption that price changes are not transmitted efficiently from one level to another and support the view that Russian retailers have more market power than farmers.

Keywords: seasonality, market power, cointegration, autoregressive distributed lags model, dairy prices

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Introduction

Price is the main tool with which different levels of the market are linked (Serra and Goodwin, 2002). Agricultural efficiency results to a large degree from the perfection of the price mechanism in the system of agents' relationships. Hence, rising food prices might provide an opportunity for agricultural development if price changes at one level (retail) were efficiently transmitted to another one (farm). However, in Russia, dairy producers express concerns about the fact that price changes are not efficiently transmitted from retailers to farmers. Price disparity has led to losses and underproduction in the rural economy. This state of play has caused redistribution of incomes from the agricultural sector to other sectors.

The phenomenon of price transmission has attracted the attention of scientists in various commodity markets. In recent years, studies have been carried out to examine price relationships between farm, wholesale and retail markets. The main focus of this research has been oriented to estimating the elasticities and speed with which shocks are transmitted between the different levels of the market chain.

Existing models that analyse vertical price transmission issues utilise several variations of a model originally introduced by Wolffram (1971) and later modified by Houck (1977). These models are based on the regression of differentiated price data and on lagged price differences where considerations can be made for the differential effects of positive and negative lagged differences. Goodwin and Holt (1999) used a vector error correction (VEC) model to evaluate monthly beef price relationships at the farm, wholesale and retail levels. They found evidence of statistically significant thresholds and asymmetries in price adjustments. Most of the literature on price transmission relies on cointegration techniques. Von Cramon-Taubadel (1998) was one of the first to incorporate the concept of cointegration into models of asymmetric price transmission. A comprehensive review of estimating and testing for asymmetric price transmission is provided in Meyer and von Cramon-Taubadel (2004).

As regards dairy products, the literature reports similar results regarding the existence of asymmetric price transmission. Serra and Goodwin (2003) identified asymmetric price relationships for sterilised milk in the Spanish dairy industry, while Lass (2005) found evidence of short-run price asymmetries in the retail milk price in the USA and observed that retail milk prices do not return to the same level following the equivalent price increases and decreases, causing an increase in the marketing margins. Stewart and Blayney (2011) have taken up the debate on asymmetric price transmission by using the threshold error correction model on milk and cheese. Bor *et al*. (2014) applied an asymmetric error correction model to monthly price data and their results suggest that there is a positive price asymmetry in the farm-retail price transmission in the Turkish milk market. Other researchers found similar asymmetries using different econometric methods: Acosta and Valdes (2013) for Panama, Falkowski (2010) in the Polish fluid milk sector and Holm *et al*. (2012) in the German milk market.

As noted above, many studies have analysed vertical price transmission using time-series econometric procedures. However, vertical price transmission in the milk market in Russia has not been investigated. In this research, vertical price transmission along the dairy supply chain in Russia (taking the case of Voronezh *Oblast* as a historically large agrarian region) is studied to gain an insight into the price interactions between the various levels of the farm-retail marketing chain.

Voronezh *Oblast* is located to the south of Moscow and has a population of approximately 2.5 million inhabitants, of whom one third live in rural areas. The Voronezh dairy sector is one of the most important, socially-significant industries. One of the major trends in the Voronezh milk market is the persistent increase in the number of dairy cattle from 2009. It is estimated that the number of milk cows in Voronezh *Oblast* rose by 3.8 per cent annually over the period 2009-2014. State support helps to maintain this trend. Within the framework of the national programme *Development of Agro-Industrial Complex*, the government subsidises and provides financial support for the renovation of existing farms and construction of new ones. Thanks to government support, investments in fresh milk production in Voronezh *Oblast* have increased significantly in recent years. Practically all the supply volume in the market comes from domestic milk producers; imported milk accounts for less than 2 per cent of supply. Milk production has increased by 15 per cent over the last five years but the productivity has dropped by 3.2 per cent (own calculations based on data from the Federal State Statistics Service of Russia). Seasonality is an important factor in milk production: the summer production volume is 2-2.5 times higher than in the low season. Voronezh raw milk producers provide about 3 per cent of total production volume in Russia. The fluid milk production is mainly from three types of milk producers: agricultural establishments (56 per cent), household farms (40.5 per cent) and private farmers (3.5 per cent).

There are problems related to price transmission and distribution of value-added between farmers and traders in the functioning of the milk supply chain. According to the National Union of Milk Producers and the Institute for Agrarian Market Studies, the farmers' share in the retail price for milk is 30-34 per cent (the suggested optimum figure in terms of incurred costs is 50 per cent) and the traders' share is 22-30 per cent (optimum: 20 per cent).

Retail sales of milk products grow annually by at least 3-5 per cent. In 2013, retail sales of dairy products in Russia increased by almost 15 per cent, including whole milk the figure was about 30 per cent. The largest retailers in the Voronezh milk market are X5 Retail Group (Russia), Tander (Russia), O'Key Group (Russia), Lenta (Russia), Auchan Group (France) and Metro Group (Germany). They control a major part the of milk retail market. The rise in retail sales of milk products is a consequence of the increasing per capita consumption level. However, per capita milk consumption has not yet reached the levels in mature economies. Increasing demand for milk is partly provided by imports but in August 2014 Russian officials introduced sanctions on dairy products and banned imports from Australia, Canada, the European Union (EU) and the USA for one year. It is envisaged that undersupply will be compensated for with imports from Belarus, Turkey and Latin American countries.

Methodology

Econometric time series and multiple regression methods were adopted for price transmission analysis. The influence of farm-gate (retail) price on retail (farm-gate) price was investigated using multiple linear regressions. The estimation of price transmission magnitude (elasticity) follows the algorithm outlined in Table 1. For a pair of prices (farm-gate and retail) for whole milk, the following steps were implemented to identify the appropriate econometric model. Depending on the price series properties, various econometric models were estimated.

Price time series are mostly non-stationary, generally leading to spurious regression. In the presence of non-stationary data, it is necessary to make them stationary by carrying out a transformation such as differencing (or detrending). Otherwise, the regression cannot be estimated correctly with ordinary least squares (OLS). Non-stationarity means presence of unit roots. In testing for the presence of unit roots, several methodological options are available. Widely used among them are 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:

$$P_t = c + \beta t + \alpha P_{t-1} + \sum_{i=1}^{k} \psi_i \Delta P_{t-i} + \varepsilon_t \qquad (1)$$

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

The PP test builds on the ADF test. While the latter uses a parametric autoregression, a great advantage of the former is that it is non-parametric. The main disadvantage of the PP test is that it works well only with large samples. It also shares some of the disadvantages of ADF tests: sensitivity to structural breaks and poor power resulting from small samples.

In a modified version of the ADF test, known as the ADF-GLS test, the time series is transformed via a generalised least squares (GLS) regression before performing the test (Elliott *et al*., 1996). The ADF-GLS test is performed analogously but on GLS-detrended data. Elliott *et al*. (1996) and later studies have shown that this test has significantly greater power than the previous versions of the ADF test.

However, it is not possible to come to a reliable conclusion about price series integration order without taking into account the seasonality in the milk markets. The approach that helps to reveal seasonal unit roots was developed by Hylleberg *et al*. (1990). The HEGY test applies to quarterly data. The seasonal unit root test for monthly data was developed by Franses (1990).

The following equation is estimated for the seasonal unit roots in monthly data:

$$\begin{aligned}\Delta_{12} P_t = {} & \pi_1 P_{1,t-1} + \pi_2 P_{2,t-1} + \pi_3 P_{3,t-1} + \pi_4 P_{3,t-2} + \pi_5 P_{4,t-2} + \\ & \pi_6 P_{4,t-2} + \pi_7 P_{5,t-1} + \pi_8 P_{5,t-2} + \pi_9 P_{6,t-1} + \pi_{10} P_{6,t-2} + \\ & \pi_{11} P_{7,t-2} + \pi_{12} P_{7,t-2} + \sum \alpha_i \Delta_{12} P_{t-i} + \varepsilon_t\end{aligned} \qquad (2)$$

where

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 autoregression (VAR) model with stationary data
		Non-stationarity	Move to step 2
2	*Cointegration* test	Exists	Estimate the Vector Error Correction (VEC) model and measure asymmetry
		No	Perform test for Granger causality and estimate vector autoregression (VAR) model using logarithmic prices in first differences

Source: own composition

$$
\begin{aligned}
P_{1,t} &= (1+L)(1+L^2)(1+L^4+L^8)P_t \\
P_{2,t} &= -(1-L)(1+L^2)(1+L^4+L^8)P_t \\
P_{3,t} &= -(1-L^2)(1+L^4+L^8)P_t \\
P_{4,t} &= -(1-L^4)(1-3^{1/2}L+L^2)(1+L^4+L^8)P_t \\
P_{5,t} &= -(1-L^4)(1+3^{1/2}L+L^2)(1+L^4+L^8)P_t \\
P_{6,t} &= -(1-L^4)(1-L^2+L^4)(1-L+L^2)P_t \\
P_{7,t} &= -(1-L^4)(1-L^2+L^4)(1+L+L^2)P_t \\
P_{8,t} &= (1-L^{12})P_t
\end{aligned} \tag{3}
$$

Where L is the lag operator in the polynomial.

Deterministic components (such as constant, trend and seasonal dummy variables) can be added to equation (2). F statistics is applied for seasonal complex roots and t statistics are applied for other roots (π_1, π_2). If the null hypothesis ($\pi=0$) cannot be rejected, it indicates the presence of seasonal unit root. The critical values are given in Franses and Hobijn (1997).

Structural breaks are often present in time series. A preliminary visual assessment of the price series in Figure 1 supports the assumption that structural breaks might be present within the period 2007-2008. To prove this, a technique developed by Zivot and Andrews (1992) was used.

Given that some price series will be non-stationary, the conventional Granger-Engle approach (Engle and Granger, 1987) which included the static following regression estimated with OLS was applied to test for co-integration:

$$\tilde{P}_{1t} = \alpha + \beta\tilde{P}_{2t} + v_t \tag{4}$$

If $\tilde{P}_{1t}$ and $\tilde{P}_{2t}$ are $I(1)$ price series, then the residuals v_t from the regression would be $I(0)$ if they are co-integrated. So, if the residuals are $I(1)$ we accept the null hypothesis of non-cointegration, otherwise, if the residuals are stationary, $I(0)$, we reject the null hypothesis and accept that $\tilde{P}_{1t}$ and $\tilde{P}_{2t}$ are co-integrated. However, the power of the Engle-Granger test is reduced if there is a structural break in the co-integrating relationship. To avoid this problem, Gregory and Hansen (1996) improved the Engle-Granger regression in order to take into account structural breaks in the intercept or in the intercept and trend.

After testing for co-integration, the Granger causality test (Granger, 1969) was applied to evaluate the possible direction of the price transmission. The starting point of the method is that P_1 variable Granger causes P_2 variable but P_2 does not Granger cause P_1.

$$P_{2t} = \sum_{i=1}^{n}\alpha_i P_{2t-i} + \sum_{j=1}^{q}\beta_j P_{1t-j} + \upsilon_t \tag{5}$$

where υ_t is the white noise, and n and q are the lag order of P_2 and P_1 variables respectively.

In this study, P_2 and P_1 are the retail and farm-gate prices, and α and β are parameters. The Granger causality test requires that the variables are stationary. In order to take into account deterministic seasonality, eleven seasonal dummies are added in the estimated regressions. In order to determine the optimum lags in the models, the Akaike Information Criterion (AIC; Akaike, 1973) and the Schwarz-Bayesian Information Criterion (BIC; Schwarz, 1978) are used. Ng and Perron (2001) proposed modified versions of AIC (mAIC) and BIC (mBIC) as a model selection criterion which are based on quasi-likelihood function.

If the price series are co-integrated, a VEC model is estimated; otherwise a vector autoregression (VAR) model for farm-gate and retail prices is built in order to investigate price dynamic relationships. The general equation of the VEC model as follows:

$$\Delta P_{2t} = \alpha + \rho(\Delta P_{2t-1} - \beta\Delta P_{1t-1}) + \delta\Delta P_{1t-1} + \theta\Delta P_{2t-1} + \varepsilon_t \tag{6}$$

where ΔP_{2t} and ΔP_{1t} are changes in retail and farm-gate prices respectively; ΔP_{2t-1} and ΔP_{1t-1} are lagged changes in retail and farm-gate prices respectively; ρ is an error correction term (speed of adjustment to long-run equilibrium); β is the long-run elasticity of price transmission; δ is the short-run elasticity of price transmission between two prices, and ε_t is the residual (white noise).

If the tests reveal non-cointegration, the VAR model can be specified and estimated. The VAR model includes two equations and can be written as follows:

$$\begin{aligned} P_{1t} = {} & \alpha_0 + \alpha_1 P_{1t-1} + \ldots\ldots + \alpha_k P_{1t-k} + \gamma_1 P_{2t-1} + \\ & \ldots\ldots + \gamma_k P_{2t-k} + \varepsilon_t \end{aligned} \tag{7}$$

$$\begin{aligned} P_{2t} = {} & \beta_0 + \beta_1 P_{2t-1} + \ldots\ldots + \beta_k P_{2t-k} + c_1 P_{1t-1} + \\ & \ldots\ldots + c_k P_{1t-k} + \varepsilon_t \end{aligned} \tag{8}$$

where P_{1t} and P_{2t} are farm-gate and retail prices, and P_{1t-k} and P_{2t-k} are lagged farm-gate and retail prices.

In the case of unidirectional Granger causality running from the farm-gate (retail) to the retail (farm-gate) price, the autoregressive distributed-lags model can be specified and the immediate and dynamic effects of one price on another estimated.

Data and empirical results

The price transmission analysis at the farm-gate and retail levels in Voronezh *Oblast* was carried out using 153 monthly observations from January 2002 to September 2014. The observations relate to nominal prices for cow whole milk per litre. The source of the data is the Federal State Statistics Service of Russia. The logarithmic transformation of monthly prices measured in RUR per litre is used. This transformation allows the results to be interpreted in percentage change terms. Analyses between prices commonly use logarithms because, with trending data, the relative error declines through time (Banerjee *et al*., 1993). Moreover, from a statistical point of view, Hamilton (1994) pointed out that the logarithmic transformation mitigates fluctuations of individual series, increasing the likelihood of stationarity after first differencing. The chain from farmers to retailers in Russia is investigated (Figures 1 and 2).

Using the methodology described above, the analysis of price series was started with the unit root tests without structural breaks. In order to select the highest number of lags for the tests, the common rule for determining P_{max}, suggested by Schwert (1989) was applied.

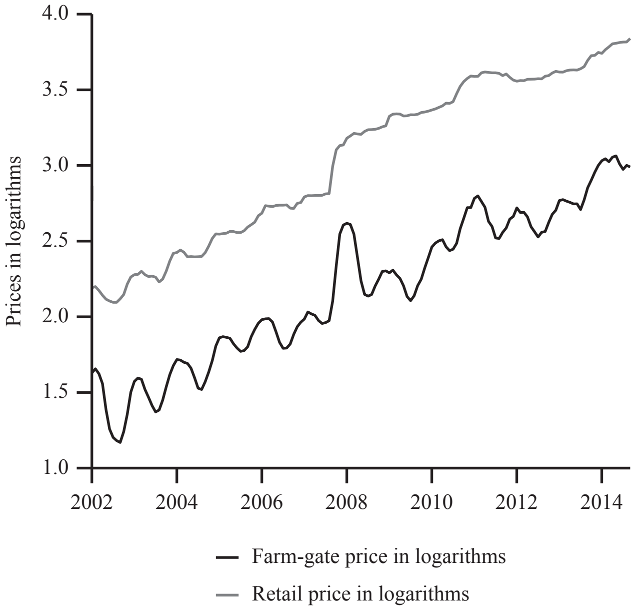

Figure 1: Price series for whole milk in logarithms in Voronezh *Oblast*, January 2002 - September 2014.

Source: own calculations based on Federal State Statistics Service of Russia data

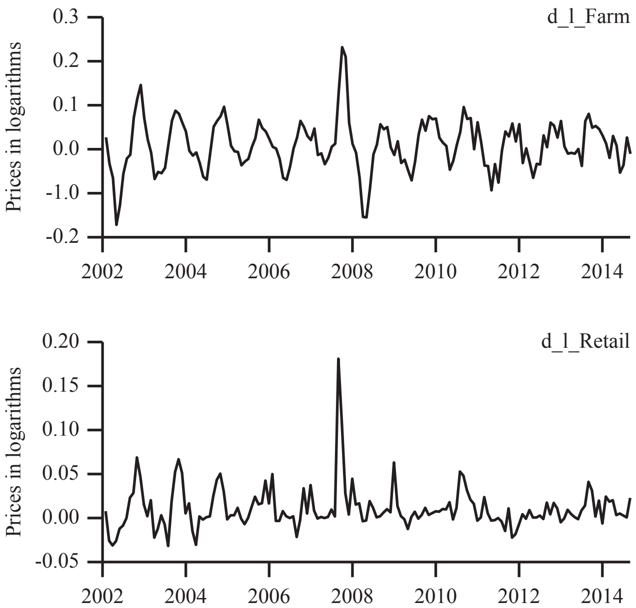

Figure 2: Price series for whole milk in logarithms (first differences), January 2002 - September 2014.

Source: own calculations based on Federal State Statistics Service of Russia data

Stationarity of the price series was checked with the conventional ADF test, ADF-GLS test, PP test, HEGY test and test with structural breaks. The number of optimal lags was determined using mBIC. The preliminary visual examination of the price series graphs provides the insight that the model for unit-root test should contain a constant and a time trend.

The null hypothesis of stationary price series in levels was rejected for all variables (Table 2). Tests based on first differences show that all the test statistics are significant at the 1 per cent level. Hence, it can be concluded that all price variables are integrated of the order one, $I(1)$. Each farm-gate and retail price series has one seasonal unit root, but not at the corresponding frequencies. So it can be concluded

Table 2: Unit root test results in levels and first differences.

Price variable (log price)	Model	ADF-GLS test			
		Lag	**Levels**	**Lag**	**First difference**
Farm-gate price	Trend and intercept	6	-1.772	1	-6.036***
	Intercept only	7	0.980	1	-5.920***
Retail price	Trend and intercept	1	-2.341	1	-6.871***
	Intercept only	1	1.660	1	-6.879***

/* null hypothesis of non-stationarity rejected at 5% and 1% of significance;
The ADF, PP, HEGY and Gregory-Hansen test results are not presented but are available from the author upon request
Source: own calculations

Table 3: Cointegration test (Engle-Granger test).

Price pair (in logarithms)	Test value	
	Intercept only	**Trend and intercept**
Whole milk (farm-retail)	-1.804 (0.628)	-2.140 (0.709)

The values in parentheses indicate p-values
Source: own calculations

Table 4: Granger causality F-test.

Null hypothesis	F-statistics, (p-value)	Conclusion
Δln*Farm_milk* does not cause Δln*Retail_milk* (lag 1)	1.050	Accept
Δln*Retail_milk* does not cause Δln*Farm_milk* (lag 1)	18.491***	Reject

Δln*Farm_milk* is the farm log-price for whole milk (in first difference); Δln*Retail_milk* is the retail log-price for whole milk (in first difference);
***/** statistically significant at the 1% and 5% levels, respectively
Source: own calculations

that there is no seasonal cointegration between them and both series are $I(1)$. Structural breaks are insignificant and are therefore not taken into account. Hence, it can be stated that the price series are $I(1)$ and that the conventional test of Engle and Granger can be run.

Within this test for co-integration the static equation (4) is first estimated with OLS and then the stationarity of the residuals of the relationship (between farm and retail prices for whole milk) is tested with the ADF test using the critical values proposed by MacKinnon (1991). ADF test statistics for the Engle-Granger test are shown in Table 3.

The null hypothesis of non-cointegration in the whole milk farm-retail chain cannot be rejected. Hence, it was found that both price pairs are not co-integrated. The VAR model can be specified and estimated in first differences. But, firstly, Granger causality F-tests of zero restrictions within the framework of VAR should be implemented. In order to estimate the possible direction of price transmission, a causality test was carried out. The appropriate lag length was selected in accordance with BIC. Seasonal dummies were added in the model. In order to avoid autocorrelation problem, heteroskedasticity and autocorrelation-consistent (HAC) standard errors within the model were computed. The direction of price transmission goes from retailers to farmers and not vice versa (Table 4).

From the findings, the ARDL (autoregressive distributed-lags) model can be specified, and immediate and dynamic effects (elasticity) of retail price on farm price for whole milk estimated (Table 5). Since the constant and time trend are statistically insignificant and also have no significant effect on the whole regression model, these variables were eliminated from the model.

Table 5: Estimation results for whole milk farm-retail chain, dependent variable ln*Farm_milk*$_t$.

Variables	Coefficient	Standard error	t-statistic	Significance (p-value)
Δln*Farm_milk*$_{t-1}$	0.41***	0.076	5.395	2.93e-07
Δln*Retail_milk*$_t$	0.312**	0.124	2.514	0.013
Δln*Retail_milk*$_{t-1}$	0.483***	0.126	3.822	0.000
S1	0.007	0.011	0.702	0.484
S2	-0.017***	0.005	-3.214	0.002
S3	-0.028***	0.006	-4.528	1.28e-05
S4	-0.025**	0.010	-2.510	0.013
S5	-0.047***	0.009	-5.353	3.55e-07
S6	-0.025***	0.007	-3.424	0.001
S7	-0.014*	0.008	-1.753	0.082
S8	0.027***	0.006	4.495	1.47e-05
S9	0.031***	0.009	3.466	0.001
S10	0.029***	0.007	3.901	0.000
S11	0.027***	0.006	4.296	3.27e-05
R^2	0.799			
Adjusted R^2	0.780			
S.E. of regression	0.028			
Residual sum of squares	0.107			
Mean dependent	0.009			
S.D. dependent	0.059			
F-statistic	38.818			1.07e-40

Estimates are given, taking into account HAC standard errors
Δln*Farm_milk* is the farm log-price for whole milk (in first difference); Δln*Retail_milk* is the retail log-price for whole milk (in first difference)
***/**/* statistically significant at the 1% 5% and 10% levels, respectively
Lag order has been selected in accordance with information criteria (BIC)

The results of the ARDL model indicate that there is a positive and significant relationship between the farm-gate and retail prices. According to the calculated price transmission elasticity, there is evidence of immediate effect that a 1 per cent increase in retail prices results in a 0.31 per cent increase in farm-gate prices. Also the dynamic long-run effect of a 1 per cent increase in retail price leads to a 1.35 per cent increase in farm-gate price.

Conclusions

This study has investigated the relationship between the farm-gate and retail prices for whole milk in the Voronezh *Oblast* of Russia. Monthly farm-gate and retail prices during the period from January 2002 to September 2014 were used in the analysis. Prices were expressed in natural logarithms to calculate percentage change. The data are integrated of order one.

Structural break tests revealed breaks but they were not significant and have not been taken into account. Vertical price transmission was evaluated in the cointegration framework, using the classical Engle-Granger approach. Visual inspection showing that price series incorporate seasonal patterns has been proved using the HEGY test. According to the findings of the research, both price series have seasonal unit roots at non-corresponding frequencies. The inclusion of seasonal dummies to reflect seasonality of price fluctuations improved the fit of the model and almost all seasonal dummies are significant. The results show that a long-run cointegration relationship does not exist between farm and retail prices, that is, they do not move together. There is evidence that change in retail price has a significant effect on farm-gate price; that is, the Granger test established unidirectional causality from retail to farm prices and not the opposite. The results on calculated price transmission elasticity revealed a short-run effect that a 1 per cent increase in retail prices leads to a 0.31 per cent increase in farm-gate prices. In addition, the dynamic long-run effect of an increase in retail price by 1 per cent causes a 1.35 per cent increase in farm-gate price. Hence, the results of this paper support the view that retailers have significant market power as highlighted by unidirectional price responses in the Russian milk market.

These findings are important for Russian policymakers in the context of the import substitution of the dairy products from the EU, USA and other countries. The results may be helpful for new policy making and support for Russian farmers. Within the framework of the new agricultural development paradigm, it is recommend to improve farmers' distribution infrastructure in order to eliminate the monopolistic power of retailers.

Further research on the topic could include the wholesale stage in the analysis to better understand price links along the dairy supply chain. Follow-up research is also needed to investigate price transmission using a wider range of advanced unit root and cointegration tests under the multiple breaks and seasonal pattern.

Acknowledgements

I would like to thank Yves Rene Surry, Professor at the Swedish University of Agricultural Sciences (Sweden), and Sebastian Hess, Professor at Kiel University (Germany) for their assistance and support during my visit to Uppsala (SLU) and in the preparation of my paper.

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