Milk Market Integration between Hungary and Poland

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Abstract

In this paper we test the retail milk price integration between two countries, Poland and Hungary. Conventional linear cointegration methods do not reveal any relationship between the two prices, therefore we apply Gonzalo and Pitakaris (2006) method to test the linear cointegration null against the threshold cointegration with an exogenous threshold variable alternative hypothesis. Our results show, that the Hungarian Forint − Polish Zloty exchange rate is econometrically an appropriate threshold variable, the linearity null is rejected, and the two alternative regimes may be characterised with different long-run equilibrium relationships. Corresponding trade data however questions the economic appropriateness of the selected threshold variable. Further research is needed to analyse the actual effective milk trade flows between the two countries, subject to exchange rate variations.

Keywords: horizontal integration, milk market, threshold cointegration

I. Introduction

Research on the spatial integration of agricultural markets is often used to test the efficiency of agricultural markets. One of the expected effects of the 2004 European Union accession is commodity price conversion on the long-run. Although horizontal integration was tested for various commodity markets around the globe, these studies mostly focus on developed economies. The question of horizontal integration might be even more interesting for transition economies, where market barriers were only recently dismantled, and market price signalling institutions are less developed. Negassa and Myers (2007) use parity bounds model to analyse policy effects and spatial market efficiency using threshold model. Goetz and von Cramon-Taubadel (2008) apply Gonzalo-Pitarakis type procedure to test between linear cointegration and threshold cointegration models. Since the linear specification is rejected, the authors propose a three step procedure to build a threshold vector error correction model, describing the price adjustments of various apples in Munich and Hamburg whole sale markets. Threshold variable is share of domestic apples in total apples marketed. In Central European Countries, Bakucs et al. (2007) and Bruemmer at al. (2009) apply Markov Switching Models to analyse horizontal integration of German and Hungarian wheat prices, and wheat prices in Ukraine respectively. For Hungary, Bakucs and Fertő (2008) use Hansen and Seo (2002) threshold cointegration test and Threshold Vector Error Correction model to test regional milk price integration.

In the post accession period, a number of producer groups perceived increased quantities of consumer milk imported mostly from Poland and Slovakia available on the supermarket shelves. Eurostat trade data (Figure 1.) shows indeed increasing post-accession trade, but the actual massive trade volume increase dates back to year 2000. Both the increased seasonality of trade and the net importer status of Hungary is clearly depicted in figure 1.

Figure 1. Polish milk imported to Hungary (IHUPOLV) and Hungarian milk exported to Poland (EHUPOLV) in EUR
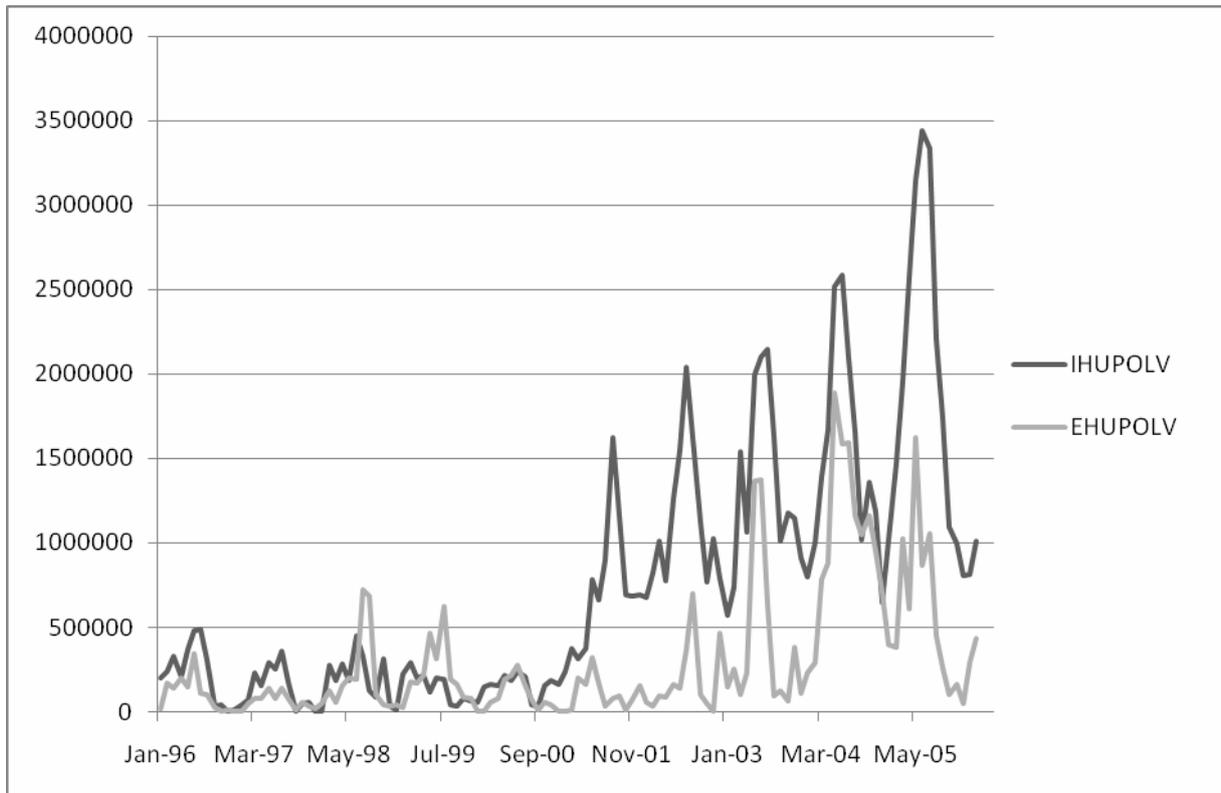
Source: EUROSTAT

The magnitude of trade relationship implies the existence of some kind of price coordination mechanism, hopefully possible to depict it econometrically. Perfectly integrated markets are usually assumed to be efficient as well. In this paper however, we simply try to analyse whether long-run price relationships underlying trade relationships exist between Poland and Hungary. Tomek and Robinson (2003), defines the two axioms of the regional price differences theory:

1. The price difference in any two regions or markets involved in trade with each other equals the transfer costs.

2. The price difference between any two regions or markets not involved in trade with each other is smaller than the transfer costs.

Let's consider, two spatially different markets, where the price of a given good in time t is P_{1t} and P_{2t} respectively. The two markets are considered integrated, if the price on market 1. equals the price on market 2. corrected with transport costs, K_t:

$$P_{1t} = P_{2t} + K_t \tag{1}$$

Trade between the two markets occurs only if $|P_{1t} - P_{2t}| > K_t$. To put it other way, the arbitrage ensures that prices of the same good traded in spatially separate markets equalise. Early studies of horizontal integration employed correlation and regression analysis. These papers usually tested some form of the Low of One Price, LOP. Consider equation (2):

$$P_{1t} = \beta_0 + \beta_1 P_{2t} \tag{2}$$

According of the *strong version* of *LOP*, prices of a given good on the spatially separated markets are equal, and they move perfectly together in time. Using the coefficients of equation (2), the necessary conditions are $\beta_0 = 0$, and $\beta_1 = 1$. In real life however, the strong version occurs only very rarely, therefore the weak version of LOP was also defined. The weak version states that only the price ratio is constant, the actual price level is different due to transport and other transfer costs. Using again the notation of equation (2), the necessary restrictions are $\beta_0 \neq 0$ and $\beta_1 = 1$.

II. Methodology

The linear Johansen cointegration procedure is based on estimating the following Vector Error Correction Model (equation 3):

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + ... + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi Z_{t-k} + u_t \tag{3}$$

,where $Z_t = [\ P^1_t, P^2_t]'$, a (2 x 1) vector containing the prices in region 1 and 2, , both I(1), Γ_1 ,....Γ_{k+1} are (2x2) vectors of the short-run parameters, Π is (2x2) matrix of the long-run parameters, u_t is the white noise stochastic term.

$$\Pi = \alpha\beta` \tag{4},$$

where matrix α represents the speed of adjustment to disequilibrium and β is a matrix which represents up to (n - 1) cointegrating relationships between the non-stationary variables. Trace and maximum Eigen-value statistics are used to test for cointegration. Once (3) is estimated we can proceed to test for weak exogeneity and then for linear restrictions on the β vector. One obvious candidate would be to test whether the elements of the vector are of the (-1, 1) form, i.e. the markets are perfectly integrated. The terms of vector α (factor loading matrix) measure the speed at which the variables adjust towards the long-run equilibrium after a price shock. The α vector of the weakly exogenous variable equals zero. To find the direction of the Granger causality between the two price series, restrictions are tested on the α vectors.

A number of studies (e.g. Barrett 2001, Fackler and Goodwin 2001, Goodwin and Piggott 2001) have questioned the appropriateness of the linear VECM models, arguing that it ignores the transaction costs that might occur. Threshold Error Correction Models (TVECM),

estimate a threshold below which the cointegration is inactive since it does not worth trading because of the low price difference. One the threshold value is exceeded, cointegration becomes active. There are various methods developed, in this paper we apply a procedure developed by Gonzalo and Pitarakis (2006). This procedure allows modelling of long-run equilibrium relationships that may change according to the level of a stationary exogenous variable. The method consists of testing the linear cointegration (equation 5) against the alternative of cointegration with threshold effects:

$$y_t = \beta' x_t + u_t \tag{5}$$

and

$$y_t = \beta' x_t + \lambda' x_t I(q_{t-d} > \gamma) + u_t . \tag{6}$$

where $x_t = x_{t-1} + v_t$; u_t and v_t being stationary disturbance terms, q_{t-d} is the lagged threshold variable and I indicator function switching between the 2 regimes. The actual test statistic is a SupLM statistic of the form:

$$LM_T(\gamma) = \frac{1}{\hat{\sigma}_0^2} u' M X_\gamma (X_\gamma' M X_\gamma)^{-1} X_\gamma' M u \tag{7}$$

where $M = I - X(X'X)^{-1}X'$.

X contains the values of x_t in (5) and X_γ contains the values of x_t, whenever the $q_t > \gamma$ condition is fulfilled, T is the sample size, $\hat{\sigma}_0^2$ is the residual variance of the linear model, and u residual term. The LM test statistic is computed across the possible values of teh exogenous threshold variable, applying a trimming parameter to ensure sufficient number of observations in each regime. The maximum value of the LM statistic (SupLM) corresponds to the test statistic, and also indicates the corresponding threshold value. Tabulated critical values are available in Andrews (1993).

Some authors (e.g. Goetz and von Cramon-Taubadel, 2008) use market shares of a given product, others transfer costs as threshold variables. In this paper we employ a different approach, using Hungarian Forint – Zloty exchange rate as threshold variable.


III. Data

Monthly time series of the log of Hungarian and Polish retail prices (figure 2.) and the Hungarian Forint-Polish Zloty exchange rate (figure 3.) between April 1997 and March 2009 were used for the empirical analysis.

Figure 2. The logarithm of the Hungarian (RPH) and Polish (RPP) retail prices of milk

RPH

RPP
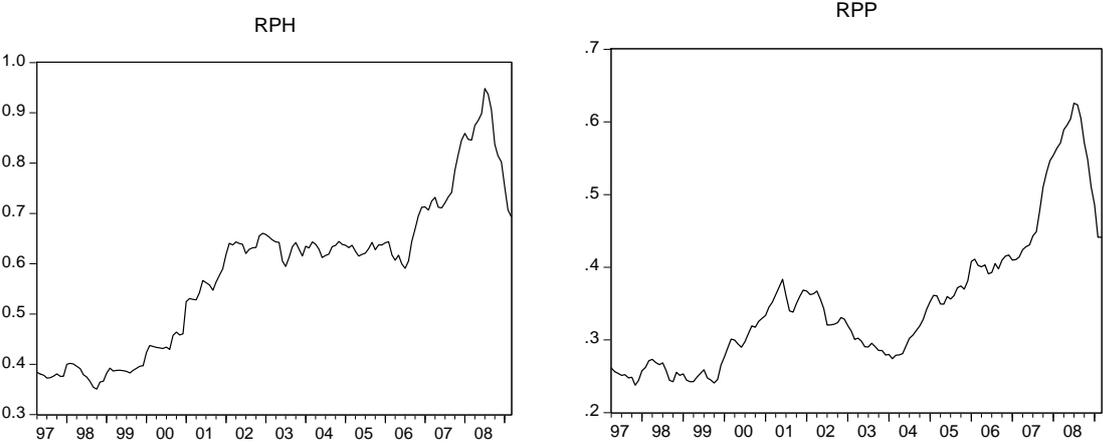
Source: Hungarian Central Statistical Agency, Polish Statistical Agency

Figure 3. The Hungarian Forint (HUF) and Polish Zloty (Zl) exchange rate

XR
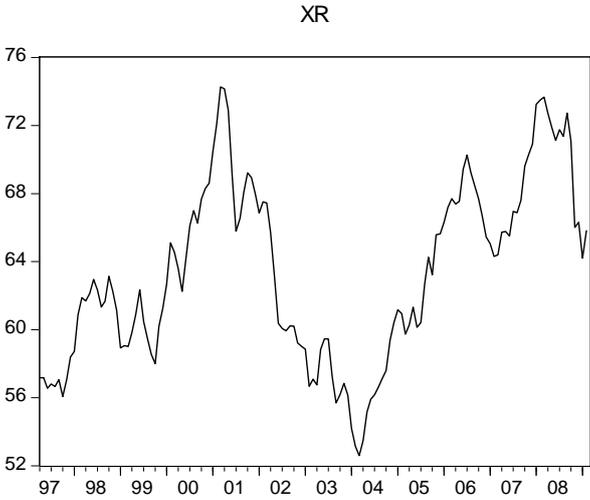
Source: Hungarian National Bank, Polish National Bank.

With potentially non-stationary data, the order of integration of endogenous variables should be tested. DF-GLS (Elliott, Rothenberg and Stock, 1996) unit root test results are presented in Tables 1 (RPH) and 2 (RPP).

Table 1. Elliott-Rothenberg-stock DF-GLS test results – Retail Price Hungary (RPH)

Specification	Level	Level with trend	First difference
Test statistic	-0.382	-2.429	-7.225
5% Crit. value	-1.943	-2.988	-1.943
10% Crit. value	-1.615	-2.698	-1.615

Table 2. Elliott-Rothenberg-stock DF-GLS test results – Retail Price Poland (RPP)

Specification	Level	Level with trend	First difference
Test statistic	-0.800	-2.331	-5.710
5% Crit. value	-1.943	-2.988	-1.943
10% Crit. value	-1.615	-2.698	-1.615

Both the Hungarian and Polish retail price of milk are integrated of order one. Conventional Johansen cointegration tests (not presented here) do not however detect cointegration between the two time series. Therefore we test the price series for threshold type integration. The SupLM statistics and the corresponding threshold values are presented on figure 4.

Figure 4. Gonzalo-Pitarakis SupLm statistics and corresponding threshold values
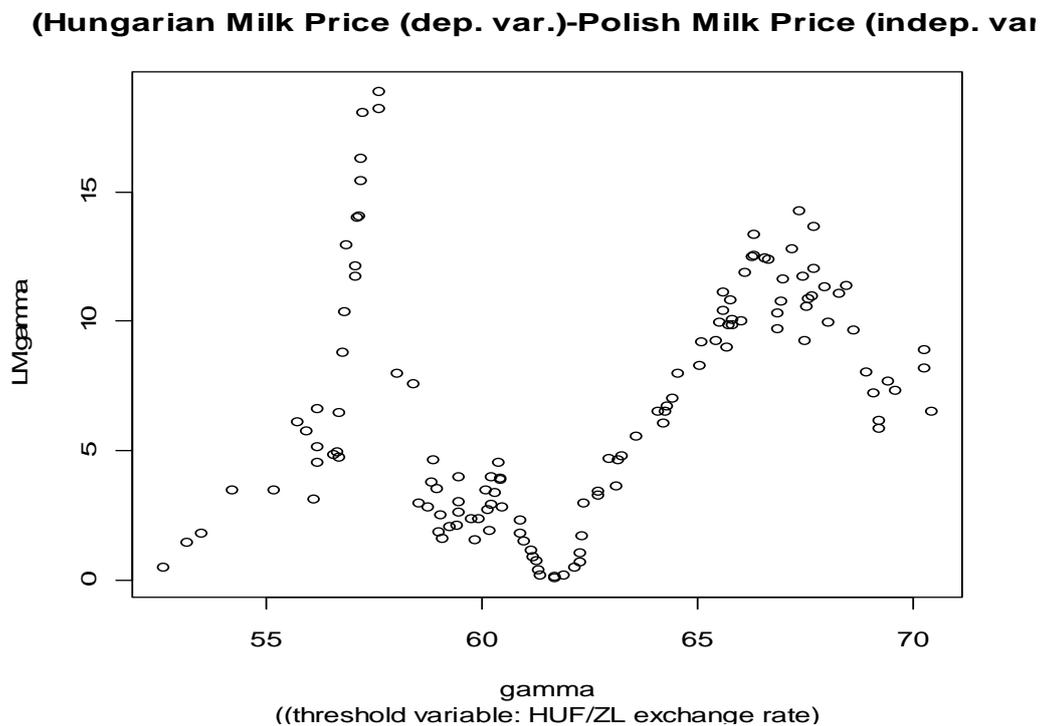
(Hungarian Milk Price (dep. var.)-Polish Milk Price (indep. var

gamma
((threshold variable: HUF/ZL exchange rate)

Source: Own calculations using Gonzalo-Pitarakis routine in R provided by Linde Götz, IAMO.

The supLM statistic reaches the maximum[1] 18.871, equivalent with a threshold of 57.59 HUF/Zl exchange rate. The SupLM is significant at 1% (critical value 16.04), rejecting the linear cointegration null hypothesis in favour of non-linear adjustments. Thus, the long-run cointegrating relationship may take different forms according to regimes. Regime I., where the HUF/Zl exchange rate is below 57.59 has 25, Regime II., with the exchange rate above the threshold has 119 observations, being the characteristic regime.

Table 3 presents the regime dependent regression coefficients, their probabilities, adjusted coefficient of determination and regression F statistic.

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[1] It is also possible to choose more than one threshold value, thus increasing the regime numbers to more than 2. Based on figure 4., one could select two threshold points (consider the 2 'peaks' of the LM statistics), the second being 14.28, significant at 5%, corresponding to a 67.38 HUF/Zl exchange rate. In this way however, there would only be 13 observations left for the 3rd regime, therefore we opted for the one threshold-two regimes model.

Table 3. Regime dependent cointegrating vectors and some statistics

Variable	Regime I. (25 obs.)		Regime II. (119 obs.)	
	Value	P - value	Value	P - value
intercept	1.792	0.003	4.174	0.027
RPP	-4.087	0.000	0.431	0.000
regression F stat	15.59	0.000	55.56	0.000
adj R^2	0.378	-	0.316	-

Source: Own calculations using Gonzalo-Pitarakis routine in R provided by Linde Götz, IAMO. Threshold point: 57.59 HUF/Zl exchange rate.

Regression F statistics are significant in both regimes, the coefficients of determination are comparable with those obtained by other studies. Intercepts and the coefficients of RPP are also highly significant. The coefficient values clearly do not satisfy the LOP (weak or strong) conditions in any of the regimes, concluding that milk price integration between Hungary and Poland is not perfect. It does however prove the existence of long-run non-linear relationships linking Polish and Hungarian milk retail prices.

Actual trade data can be used to get an insight on the economic meaningfulness of the estimated regimes. Trade volume is generally higher in the second regime, the share of Polish imports in total bi-lateral trade is 56.5% in the first regime, and 76.7% in the second regime. This is against our expectations, since it suggests increasing imports in total trade whenever national currency depreciates, and vice-versa for appreciation, this being against the economic logic. This in turn might rise questions about the appropriateness of the exchange rate as a threshold variable.

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