Spatial product market integration between two small, open neighbouring economies

Zoltán Bakucs

Institute of Economics – Centre for Economic and Regional Studies, Hungarian Academy of Sciences, 1112 Budapest, Budaörsi u ́t 45, Hungary and Corvinus University of Budapest, 1093 Budapest, Fővám tér 8, Hungary. E-mail: zoltan.bakucs@krtk.mta.hu

Stefan Bojnec

Faculty of Management, University of Primorska, 6104 Koper, Cankarjeva 5, Slovenia. E-mail: stefan.bojnec@fm-kp.si; stefan.bojnec@siol.net

Imre Fertő

Institute of Economics – Centre for Economic and Regional Studies, Hungarian Academy of Sciences, 1112 Budapest, Budaörsi u ́t 45, Hungary and Corvinus University of Budapest, 1093 Budapest, Fővám tér 8, Hungary. E-mail: imre.ferto@krtk.mta.hu

ABSTRACT

This paper contributes with an in-depth investigation of spatial wheat producer market integration between two neighbouring countries, net exporting Hungary and net importing Slovenia. One of the key features of this spatial producer price transmission is the very important role Hungary plays in supplying wheat to Slovenia. Using monthly data from January 2000 to April 2011, spatial price transmission is analysed from a number of perspectives, using a wealth of econometric techniques in order to shed light upon the degree of integration, adjustment asymmetries and the role of market share upon price transmission. Empirical results rejected the validity of the Law of One Price, identified Hungary as the price-leading market, confirmed competitive symmetric price adjustment, and emphasized the role of Hungarian market share in total Slovenian imports with respect to wheat price transmission elasticity. The results imply that wheat producer prices in both of the analysed countries tend to be internationally determined. Hungarian exports are unable to price discriminate in wheat trade between the two neighbouring markets.

Keywords: Spatial price transmission; Law of One Price; Price-leading market; Cointegration; Hungary and Slovenia

JEL classification: F15, C12, C21, D43, Q13

1. INTRODUCTION

One of the most important targets of the European Union's (EU) Common Agricultural Policy (CAP) is to facilitate the spatial integration of agricultural markets within the individual member states as well as within the EU. On a spatially integrated market, the price of a product and price information should freely be transmitted between member states to attain an integrated and efficient market.

Although global wheat market is increasingly concentrated in last decades, the role of regional wheat trade also plays an important role. The intra EU wheat trade is dynamically expanding, amounting to 80% (2011 data) of the total EU wheat trade (Eurostat, 2012). Consequently, the validity of the *Law of One Price (LOP)* within the EU is an important research and agribusiness policy issue revealing the efficiency of the EU's spatial market integration.

This study provides a quantitative analysis of the wheat producer pricing behaviour on two neighbouring wheat markets in the EU. Namely, the paper aims to investigate spatial wheat producer market integration between Hungary and Slovenia, which are two small open neighbouring Central European economies. These two countries are not competing wheat exporters. Hungary is a major wheat exporting country in the group of the ten New Member States (NMS-10) in the 2004 EU enlargement, whilst Slovenia is traditionally net wheat importer. Wheat producer price difference between the two neighbouring countries is expected to lead to price arbitrage from the global wheat market to close the gap between the two prices. Therefore, we expect the co-movement of prices under regional and global wheat market price arbitrage, which prevents price discrimination on long-run.

The wheat trading relations between Hungary and Slovenia have been important before and after the 2004 EU accession. Wheat producer price levels are important for the agricultural markets on both the revenue (wheat sales) and the cost (animal fodder) side of many farms' balance sheets. On the demand side, wheat is important for industrial and bioenergy uses, for concentrates, as well as for human nutrition. Hence, the development of the wheat producer price transmission between Hungary and Slovenia is an important topic for spatial market integration analysis.

For a small open economy, such as Hungary or Slovenia, market efficiency, and market information flow have at least two important business and economic policy consequences. The first one is the transmission of prices by some actors of the food supply chain either vertically or spatially. This issue of international market competitiveness is quite relevant for Hungary and Slovenia, considering the structure of their agro-food markets, which has developed from regionally monopolized markets during the past communist regimes. The second consequence relates to the agricultural support system completing the CAP in the NMS-10, including Hungary and Slovenia. This paper focuses on the first consequence and topic for global business studies, by testing for price transmission between Slovenian and Hungarian wheat producer prices.

The paper is organized as follows. We start by presenting a background on literature review on agri-food business studies focusing on spatial price transmission in wheat markets. Then, we discuss spatial product market integration, followed by an introduction to the methodology. The core of this paper is the fifth part, the empirical analysis focusing on various aspects of market integration. Following the discussion of results, the final section derives main conclusions.

2. BACKGROUND ON SPATIAL PRICE TRANSMISSION ON WHEAT MARKET

The previous literature on spatial price transmission in wheat markets has mainly focused on US-Canada relationships, main global exporters and international markets (e.g. Pick, & Park, 1991; Pick, & Carter, 1994; Mainardi, 2001; Mohanty & Langley, 2003; Carew, & Florkowski, 2003; Bessler, Yang, & Wongcharupan, 2003; Ghoshray, 2002, 2007; Tun-Hsiang, Bessler, & Fuller, 2007; Jin, & Miljkovic, 2008). Bukenya and Labys (2005) argue that in a spite of improvements in communications and globalization, the empirical results do not support the convergence of wheat commodity prices on world commodity markets in spatially dispersed markets during the period 1930-1998, but a pattern of fluctuating divergences.

However, there are just a few studies on the European wheat markets (e.g. Ejrnæs & Persson, 2000; Thompson, Sul, & Bohl, 2002; Dawson, Sanjuan, & White, 2006; Brosig et al., 2011; Eryigit and Karaman, 2011). Pall et al. (2013) provide a quantitative analysis of the pricing behaviour of Russian wheat exporters. And only a few studies concentrate on relationships between EU-15 and a Central and Eastern European (CEE) country (e.g. Bakucs et al., 2012).

To the best of our knowledge, there is no published research focusing on spatial product market integration of cereal prices between two CEE countries, and particularly not between Hungary and Slovenia. Most of wheat trade between these two countries is soft wheat as a rather homogenous product in bilateral trade and in domestic wheat production. In a spite of possible deficits in the development of market institutions and market inefficiencies, the *LOP* and symmetric price transmission might hold as global price arbitrage can prevent possible price discrimination. Therefore, the evolution of spatial price transmission and functioning of regional wheat markets is perhaps of even more interest in transition and emerging market economies than in developed ones.

This paper adds to the existing literature of spatial product market integration by the detailed analyses of horizontal price transmission between Hungarian and Slovenian wheat markets at the producer level, using monthly wheat producer price data from January 2000 to April 2011. We employ a battery of econometric techniques, starting with a linear Vector Error Correction model (VEC), followed by two non-linear cointegration frameworks, each shedding light into different aspects of product market integration.

3. SPATIAL PRODUCT MARKET INTEGRATION

Research on the spatial integration of agricultural markets is often used to test the efficiency of agricultural markets. Perfectly integrated markets are usually assumed to be efficient. Tomek and Robinson (2003) define the two axioms of the international price differences theory:

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

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

Transfer costs are equal to transportation costs between markets and various other handling costs.

Let's consider, two spatially different markets, where the price of a given good on market 1 in time *t* is P_{1t} and on market 2 in time *t* is P_{2t} respectively. The two markets are considered integrated, if the price on market 1 equals the price on market 2 corrected with transportation and other handling 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 international markets equalize. Empirical literature usually tests the validity of the `*LOP* by considering the following equation, with prices expressed in logarithms[1]:

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

According of the *strong version* of *LOP*, prices of a given good on the spatially separated international 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* of *LOP* occurs only very rarely, therefore the *weak version* of *LOP* was also defined. The *weak version* of *LOP* states that only the price ratio is constant, the actual price level is

[1] Using logged data facilitates the interpretation of results since estimated coefficients are price transmission elasticities (i.e., 1% change in the independent variable induces β_1% change in the dependent variable.

different due to transportation and other handling or transfer costs. Using again the notation of equation (2), the necessary restrictions are $\beta_0 \neq 0$ and $\beta_1 = 1$.

With the evolution of time series econometrics, recent papers (e.g. Balcombe, Bailey, & Brooks, 2007) test a more general (wider) notion of horizontal integration of spatially separated international markets. In this case the long-run co-movement of prices is analyzed. Therefore, we test the following hypothesis (H) for the validity of the *strong* or *weak versions* of *LOP*:

H_1: The law of one price holds for wheat producer prices between the neighbouring wheat net exporting country Hungary and wheat net importing country Slovenia.

Although there are some Slovenian wheat exports to Hungary, these quantities are irrelevant for Hungarian wheat market compared to the much larger Hungarian exports to Slovenia. Thus it is plausible to expect Hungarian prices to cause Slovenian ones, and not vice versa, formally tested by H_2:

H_2: On long-run, Hungarian wheat producer prices are determining the Slovenian wheat producer prices.

World wheat markets are considered to be oligopolistic, where the US, Canada, Australia and the EU account for 90% of international trade (Ghoshray, 2010). The two small open economy players, i.e. Hungary and Slovenia, both EU member states, are expected to respond to international market conditions, but be unable to influence them. This suggests that they are more likely to be price takers with respect to international wheat trade prices and/or prices at wheat stock exchanges. Considering the relatively small Hungarian and Slovenian market shares in global wheat production and trade (FAO, 2012), it is unlikely that their traders may be able to exercise market power. Therefore, it is also unlikely that they may be able to impose an asymmetric adjustment to the long-run bilateral price equilibrium as response to negative or positive short-run price deviations (shocks). Therefore, we test the following H_3 on competitive international wheat market conditions:

H_3: Given the competitive international wheat market conditions, adjustment between Hungarian and Slovenian wheat producer prices is symmetric, i.e. negative and positive price deviations from long-run equilibrium are transmitted with equal speed.

And finally, as a number of empirical studies has shown (e.g. Götz & von Cramon-Taubadel, 2008; Stephens et al., 2008) the *LOP* might hold or not[2], depending on trade volumes, or more specifically of the market share of products from country A to country B. Thus, we test the following hypothesis:

H_4: The magnitude of price elasticity (β_1 in equation (2)) depends on the share of wheat imported from Hungary in total Slovenian wheat imports.

[2] A more relaxed assumption states that long-run price elasticities might be different depending on trade volumes (shares).

4. METHODOLOGY

Our three stage empirical strategy follows the set hypotheses outlined in the previous section, and is nested within the non-stationary time series framework. After determining the order of integration of the series, we employ system and single equation cointegration tests, followed by the estimation of VEC models (where appropriate)[3]. We start with the most commonly used linear cointegration tests, to check whether long-run co-movement of prices exists and whether strong or weak versions of *LOP* holds. Second, we assess whether there are or not asymmetries with respect to the speed of adjustment using the Enders and Siklos (2001) framework. And finally, we employ the Gonzalo and Pitarakis (2006) approach which allows us to derive conclusions whether the intensity of wheat trade between Hungary and Slovenia has implications upon the horizontal price transmission mechanism.

The widely used Johansen (1988) maximum likelihood cointegration test is nested within the more general Johansen, Mosconi, and Nielsen (2000) test, allowing for up to two breaks in the long-run equation. The procedure estimates the following model:

$$\Delta Y_t = \alpha \begin{pmatrix} \beta \\ \mu \end{pmatrix}' \begin{pmatrix} Y_{t-1} \\ tE_t \end{pmatrix} + \gamma E_t + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + u_t \tag{3}$$

where Y_t is a vector of non-stationary variables (in our case the Hungarian and Slovenian wheat producer prices), p is the lag number, $E_t = (E_{1t}\ E_{2t} \ldots E_{qt})'$ is a matrix of q dummy variables, where $E_{j,t} = 1$ if observation t belongs to the j^{th} period and 0 otherwise, $D_{j,t-i}$ is an impulse dummy that equals 1 if observation t is the i^{th} observation of the j^{th} period, meant to render the corresponding residuals to zero. Γ_i and $K_{j,i}$ are short run matrices, α is the speeds of adjustment parameter matrix, β are the long run cointegration coefficients and μ are the long run drift parame-

[3] The widely used, linear, Johansen methodology is a multi-equation Vector Error Correction model couched approach, especially suitable if multiple cointegrating relationships are likely. The currently applied non-linear (e.g. endogenous structural break, threshold or switching model) cointegrating tests are however mostly single-equation approaches. The Granger Representation Theorem (Engle and Granger, 1987) however states that for every cointegrating relationship exists a Vector Error Correction representation, allowing the estimation of a VEC model irrespective whether the cointegration relationship originates from single or multi-equation models.

ters. The u_t residuals are supposed to be independently and identically distributed with zero mean and symmetric and positive definite variance-covariance matrix Ω. Restrictions on the model on *LOP* and weak exogeneity of the Hungarian wheat producer prices (H_1 and H_2) can be tested using likelihood ratio tests. Equation (3) may be re-written as:

$$\Delta Y_t = \alpha \varepsilon_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + u_t \tag{4}$$

where ε_t is the stationary error correction term, in fact the residuals of the long-run cointegration equation. Equations (3) and (4) assume that the adjustment process is symmetric, i.e. positive and negative price deviations induce the same adjustment (H_3).

Enders and Siklos (2001) propose a more flexible threshold cointegration test, allowing different adjustment speeds with respect to positive and negative price changes. The respected test equation has the following form:

$$\Delta \varepsilon_t = I_t \rho_1 \varepsilon_{t-1} + (1 - I_t) \rho_2 \varepsilon_{t-1} + \gamma_1 \Delta \varepsilon_{t-1} + ... + \gamma_p \Delta \varepsilon_{t-1} + v_t \tag{5}$$

where ε_{t-1} are the lagged error correction terms, $\Delta\varepsilon_{t-i}$ lags of the dependent variable, to account for residual serial autocorrelation, and v_t independently and identically distributed residuals, and finally, I_t the Heaviside function with the following properties:

$$I_t = \begin{cases} 1, & if \quad \varepsilon_{t-1} \geq \tau \\ 0 & if \quad \varepsilon_{t-1} < \tau \end{cases} \tag{6}$$

where τ is the threshold value. The test regression is appropriate (convergent) if ρ_1 and ρ_2 coefficients are negative. Cointegration is tested by the so called T_{max} and Φ statistics, both with non-standard distribution. More importantly, the null hypothesis of symmetrical adjustment ($\rho_1 = \rho_2$) against the alternative of asymmetric adjustment is tested using an F-test. Should adjustment asymmetries be incorporated into a VEC model, equation (4) displays the following form:

$$\Delta Y_t = \alpha_1 I_t \varepsilon_{t-1} + \alpha_2 (1 - I_t) \varepsilon_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + v_t \tag{7}$$

Whilst the Enders and Siklos (2001) method, allows for non-linear short-run adjustment process, the long-run cointegrating relationship is unique and linear. Gonzalo and Pitarakis (2006) propose the following cointegrating regression with threshold non-linearity, where not only the short-run process, but also cointegrating equilibrium relationship can vary according to the dynamics imposed by a stationary exogenous variable. Let y_t and x_t be the Slovenian and Hungarian wheat prices, and q_{t-d} with d≥1 a stationary exogenous variable defining the regimes:

$$y_t = \beta^{'} x_t + \lambda^{'} x_t I(q_{t-d} > \lambda) + u_t \tag{8}$$

$$x_t = x_{t-1} + v_t \tag{9}$$

After estimation, the following sup LM statistic (equation (11)) is used to test the null hypothesis of linear cointegration (10) against the alternative of cointegration with threshold adjustments (8):

$$y_t = \beta^{'} x_t + u_t \tag{10}$$

$$LM_T(\gamma) = \frac{1}{\sigma_O^2} u^{'} M X_\gamma (X_\gamma^{'} M X_\gamma)^{-1} X_\gamma^{'} M u \tag{11}$$

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

and X stacks x_t corresponding to the linear, and X_γ stacks $\mathbf{x_t}I(q_{t-d}>\gamma)$ for the non-linear model, u_t and v_t are scalar and stationary disturbance terms, $I(q_{t-d}>\gamma)$ is the indicator function taking the value of 1 when $q_{t-d}>\gamma$ and 0 otherwise. Depending on the test result, a linear or regime dependent VEC model may be estimated.

5. EMPIRICAL ANALYSIS

5.1. Data

Monthly wheat producer prices for Hungary and Slovenia from January 2000 to April 2011 are used (Table 1). Data source for Slovenia is the SI-STAT of the Statistical Office of the Republic of Slovenia (SORS, 2011). Data are reported in Euro, which replaced the Slovenian Tolar (SIT, the former Slovenian national currency) after 1st January 2007. Price data for Hungary are obtained from the Hungarian Central Statistical Office. Hungarian wheat producer prices were reported in Hungarian Forint (HUF) per kg, however, in order to facilitate the analysis, and to avoid inflation induced trends, data was transformed into Euro per kg using monthly average Euro per HUF exchange rate[4]. In addition, wheat trade data between Hungary and Slovenia were obtained from the Eurostat Comext trade database (Eurostat, 2012).

At the first glance, as shown on Figure 1, wheat producer price developments in Hungary and Slovenia follow rather similar patterns over time that might suggest the validity of H_1. Yet, these wheat producer price level patterns might indicate that Slovenia, as a traditionally net wheat importer, largely follows producer price developments of the net wheat exporter Hungary. This might imply the validity of H_2, i.e., Slovenian wheat producer prices are determined by the Hungarian wheat producer prices.

The impact of the EU accession on May 1st, 2004 is not visible on wheat producer price series. Interesting pre- and post-accession differences may be noted however on the trade share (Figure 2). Because of the already liberalized pre-accession regional trade relations between Hungary and Slovenia due to the Central European Free Trade Agreement, the importance of Hungarian wheat exports in total Slovenian imports are equally important in the pre- and post-accession period. During the pre-accession period, the volatility of Hungarian export share in total wheat imported to Slovenia was much higher (the standard deviation in the pre-accession period is 0.36 against 0.15 in the post-accession period).

- - insert Table 1 here –insert Figure 1 here –

- insert Figure 2 here –

5.2. Quantitative results

[4] Note, that Euro adoption by Slovenia, and the fact that Hungarian prices reported in HUF were transformed into Euro, should not affect results, since international (European) wheat trade uses prices denominated in Euros.

The empirical analysis of wheat market integration between Hungary and Slovenia is done in four steps. First, we assess the time series properties of data, followed in a second step by the estimation of a linear VEC model, capable to assess the validity of the *LOP* (H_1) and the causality relations between Hungarian and Slovenian wheat producer prices (H_2). In a third step, we estimate a non-linear, threshold cointegration model, testing whether adjustments to departures from long-run equilibrium are equal in case of negative and positive price deviations (H_3). Finally, we assess the impact of wheat trade volumes on wheat producer price market integration, estimating a wheat trade share dependent non-linear model, again suitable for *LOP* tests (H_4).

5.2.1. Time series data properties

The logged Hungarian (PWHU) and Slovenian (PWSVN) wheat producer prices, as well as the share of wheat imports from Hungary in total Slovenian wheat imports (SHARE_HU) are tested for unit roots[5]. Table 2 reports the Augmented Dickey-Fuller unit root test results for the time series and their first differences (second part of Table 2). At 5% level of significance, the null hypothesis of unit root cannot be rejected for the price series in levels, regardless of the deterministic specifications (constant or constant and trend). The first differences are however stationary, thus both time series are integrated of order 1, i.e. contain one unit root. On the other hand, the SHARE_HU variable proved to be stationary, rejecting the unit root null for both the time series in levels and its first difference at even 1% level of significance.

- Insert Table 2 here -

5.2.2. The Law of One Price and the causality relations

[5] There are a large number of unit root tests in the empirical literature. Unit root tests in the presence of structural breaks, to control for possible EU accession or Euro adoption (Slovenia) effects were also applied. Results, however, confirmed ADF outcomes, with break dummies for above periods not being statistically significant.

Since the time series for the Hungarian and Slovenian wheat producer prices, respectively, were found to be non-stationary, the cointegration framework is generally suitable for further analysis.

First, we proceed with linear cointegration, followed by the estimation of a linear VEC model. Both the Johansen et al. (2000) trace, and the Saikkonen and Lütkepohl (2000) tests reject the no cointegration (0 cointegration vectors) hypothesis at 6% (or 1%) level of significance, however the null hypothesis of a single cointegration vector cannot be rejected (p=0.6 and 0.8 respectively) concluding that the Hungarian and Slovenian wheat producer price series are cointegrated with one cointegrating vector (Table 3).

- insert Table 3 here –

Using the cointegration test results, a linear VEC model (equation 3) is estimated in the next step. Information criteria selected one lagged autoregressive short-run term. Table 4 presents the results.

- insert Table 4 here –

LM tests[6] up to the 4th lags, and Portmanteau autocorrelation tests up to the 12th lags emphasize that VEC model is well specified, residuals do not suffer from serial autocorrelation at 5% level of significance. The Jarque-Bera statistics, however, reject the residual normal distribution hypothesis at conventional levels. Residual non-normality implies that the test results must be interpreted with care, although asymptotic results do hold for a wider class of distributions (von Cramon-Taubadel, 1998).

Thus, the long-run relationship between Hungarian and Slovenian wheat producer prices in equation (12) with t-statistics in parentheses is:

[6] Diagnostic tests are not presented here to save space. They are available from the authors upon request. Largest significance level recorded for LM tests was p=0.098 (3 lags), for Portmanteau tests p=0.376 (7 lags).

$$PWSVN = - 0.79 + 0.558*PWHU + u_t \qquad (12)$$

$$(4.03) \qquad (-6.13).$$

Thus 1% increase in Hungarian wheat producer prices induces a 0.55% increase in Slovenian wheat producer prices. The highly significant coefficients do not seem to support neither the *strong* nor *weak* versions of the *LOP*. The formal likelihood ratio test of the null hypothesis $\beta_{PWHU} = 1$, results a $\chi^2(1)=8.79$ (p=0.003) significant at 1%, thus rejecting the H_1 null. The speed of adjustment coefficient – α in equation (3) and (4) – shows the magnitude of system adjustment towards the long-run equilibrium after a shock to the system. It is highly significant, with the right sign in the Slovenian equation ($\alpha_{\Delta PWSVN}$ = -0.358 with a Student t-statistic =-4.109), and not significant in the Hungarian equation ($\alpha_{\Delta PWHU}$ = -0.06 with a Student t-statistic = -0.962) suggesting as expected, the weak exogeneity of Hungarian prices. Formal likelihood ratio testing the null hypothesis of weakly exogenous Hungarian prices, and thus, long-run Granger causality (Granger, 1981; Engle and Granger, 1987) running from Hungarian towards Slovenian prices, result $\chi^2(1)=0.786$ with a significance level of p=0.73. This result implies that as *a priori* expected, Hungarian wheat producer prices determine Slovenian wheat producer prices, not rejecting H_2 null.

5.2.3. Adjustments of price deviations from long-run equilibrium

The linear model, however, is not capable to offer information with respect to the symmetry or asymmetry of price adjustments. To assess whether Slovenian wheat producer price adjustment to long-run equilibrium is responding in a similar fashion to Hungarian wheat producer price increases or decreases, next we employ the Enders and Siklos (2001) threshold cointegration test. The threshold value is set to zero[7], thus the obtained ρ_1 and ρ_2 coefficients (above and be-

[7] It is possible to use alternative threshold values, based on *a priori* information, or endogenously searched by minimizing the error sum of squares of the fitted model. Search procedures result threshold values close to zero, and do not affect the test conclusions. Thus, they are not discussed here.

low threshold, see equation (5)) reflect adjustments of the positive and negative price deviations from long-run equilibrium. The test results are presented in Table 5.

- insert Table 5 here -

The empirical results in Table 5 confirmed the following main empirical facts and findings. First, the model is convergent as the estimated above and below threshold coefficients are significant, with the right (negative) sign. Second, both the *T-max* and *Φ* statistics are above their simulated critical values, rejecting the null of no cointegration, reinforcing the Johansen Mosconi, and Nielsen (2000) and Saikonnen and Lütkepohl (2000) cointegration results. Third, the estimated below threshold coefficient is higher (in absolute value) than the above threshold coefficient, suggesting that price increases with respect to the long-run equilibrium persist longer (phase out slower) than negative deviations. These results would imply asymmetric adjustment, but a formal F test of coefficient of equality results a test statistic of $F^{*}=0.406$ against the simulated critical value of $F_c=2.46$. This statistical fact points in favour of not rejecting the coefficient equality null hypothesis. Thus, we may conclude that adjustments to long-run equilibrium in the Hungarian and Slovenian wheat price relations are symmetric, and do not reject H_3 null. This result could be somewhat expected, bearing in mind that both Hungary and Slovenia are small open economies, with developed infrastructure and actors in a competitive international wheat producer market. Producers in both countries are price takers in the international wheat markets. The neighbouring Hungary is competitive in the Slovenian wheat market due to the geographical closeness, and thus possibly lower transportation costs. On the other hand, Slovenia can also easily buy wheat from other countries in the region, or by imports from the rest of the world, through the Slovenian port Koper. Therefore, the empirical results confirm mutual interests for competitive spatial international wheat producer price integration and rather smooth price transmission. Thus the asymmetric VEC model estimated according to equation (7) is reduced to the equation (4), and the VEC model results in Table 4.

5.2.4. The importance of bilateral wheat trade

We proceed to test the importance of trade flows between Hungary and Slovenia. More exactly, whether the wheat imported from Hungary in total wheat imported by Slovenia plays any role in horizontal spatial wheat producer price market integration and *LOP*. The choice of the stationary (Table 2) threshold variable, SHARE_HU, reflects the importance of spatial market price integration of Slovenia to the Hungarian wheat producer market (see Figure 2 and Table 1). It seems plausible, that long-run Hungarian-Slovenian wheat producer price relationships may display different patterns depending on the actual quantity/share of Hungarian wheat imported on the Slovenian market. Thus, the variable SHARE_HU is used as a threshold variable for the Gonzalo and Pitarakis (2006) cointegration test[8]. The estimated LM_γ statistics (equation (11)) and the threshold values are presented on Figure 3.

- insert Figure 3 here -

The maximum LM_γ statistic is 12.088, significant at 1%, and corresponds to a threshold value of 0.5719. Thus two regimes with distinct long-run parameters are identified. The first regime contains 26 observations, and is active when the share of Slovenian imports from Hungary in total Slovenian imports is less than 57%. The second, larger regime pools 110 observations, when the wheat import share from Hungary in total Slovenian wheat imports is more than 57%. The estimated regime dependent VEC model is presented in Table 6.

- insert Table 6 here -

[8] For the empirical test, we use a program written in R language, courtesy to Linde Götz, Leibnitz Institute of Agricultural Development in Central and Eastern Europe, Halle (Saale), Germany.

The non-linear VEC model in Table 6 is composed of 2 regimes, displaying coefficients with correct signs. The smaller (first) regime is less well specified with some evidence of residual serial autocorrelation at smaller lags. For the regime with most observations (second), there is no trace of residual serial autocorrelation up to the 4th lag[9], and most of the autoregressive terms are significant. The residual normality assumption is only rejected for one of the four equations, namely for the Hungarian wheat producer price equation in the second regime[10]. The long-run relationship between Hungarian and Slovenian wheat producer prices, conditioned on the magnitude of SHARE_HU variable (for constant and slope coefficients significance, see Table 6), is:

$$PWSVN = \begin{cases} -1.015 \quad +0.497PWHU \quad if \quad SHARE\_HU < 57.2\% \\ -0.572 \quad +0.650PWHU \quad if \quad SHARE\_HU >= 57.2\% \end{cases}$$

Thus, depending on the share of Slovenian wheat imports from Hungary (below or above 57.2%), 1% increase in Hungarian wheat producer prices induce a 50% or 65% price increase in Slovenian wheat producer prices. Similarly to the linear model, the highly significant coefficients do not seem to support neither the *strong* or *weak* versions of the *LOP*. The formal likelihood ratio test of the null hypothesis $\beta_{PWHU} = 1$, result $\chi^2(1) = 67.29$ and $\chi^2(1) = 76.30$ for the first and second regime respectively, both significant at 1%, thus again rejecting the H_1 null. The speed of adjustment coefficients are significant (p=0.04 and 0.01 respectively), with the right sign for the Slovenian equations ($\alpha_{\Delta PWSVN}$ = -0.57 and -0.37), and statistically not different from 0 for the Hungarian equations (likelihood ratio test results for the two regimes are $\chi^2(1)$ = 1.57 and $\chi^2(1)$ = 1.15 respectively, none significant) reinforcing the previous results with respect to causality running from Hungarian towards Slovenian wheat producer prices (H_2). As

[9] For the first regime, the most significant test statistic recorded for the Slovenian and Hungarian equations was at lag 1 (p=0.08) and significant for lag 2 (p=0.002). For the better specified second regime, the most significant test statistic for the system was for lag 3 (p= 0.19).

[10] Residual normality is desired for VEC models. However, asymptotic results do hold for a wider class of distribution (von Cramon-Taubadel, 1998).

with the linear model, this result implies that, as *a priori* expected, Hungarian wheat producer prices determine Slovenian wheat producer prices, not rejecting H_2 null.

To sum up, Gonzalo and Pitarakis (2006) exogenous threshold driven cointegration confirmed previous linear and non-linear results, with respect to the rejection of either strong or weak versions of the *LOP*. In addition, it detected two trade-share dependent long-run price elasticity relationships, albeit with rather small difference (49% and 65% in the first and second regimes, respectively). Table 7 presents some descriptive statistics of the Hungarian share in total monthly Slovenian wheat imports (SHARE_HU) in the two regimes.

- insert Table 7 here -

The larger regime (price elasticity is 0.65), with 110 observations |(108 after adjustments for autoregressive terms) may be considered the ‘normal regime’, when on average 82% of Slovenian wheat imports originate from Hungary. Actually, in some months, all Slovenian imports originate in Hungary. In addition, the Slovenian Adriatic coast port Koper is also used for Hungarian wheat exports, and this port is also used for Slovenian wheat imports from the rest of the world. As a possible consequence, wheat producer prices in additional to bilateral wheat trading relations are determined by the world or international wheat prices, which in a case of Hungarian wheat export, this is f.o.b. price, while in a case of Slovenian wheat import, this is c.i.f. price. As a result, these international wheat price relations determine also bilateral wheat producer prices in Hungary and Slovenia. In contrast, the small regime (price elasticity equals 0.50) is when Hungary plays a smaller role in Slovenian wheat imports, averaging 25.6% of total imports. The maximum in this regime equals the threshold, whilst in a few cases there are no imports at all from Hungary. Another interesting statistic differentiating the two regimes is the standard deviation/mean ratio (relative standard deviation). This statistic equals 0.77 in the first regime, but just 0.15 in the larger regime suggesting not just higher volatility of traded quantities, but also that the small regime collects the extremes and is characterized by sudden changes.

6. DISCUSSION AND AGRIBUSINESS IMPLICATIONS

To sum up, very robust results were obtained. We employed three different methodologies for an in-depth assessment of the wheat producer price relationship between Hungary and Slovenia, its determinants and more specifically whether the *LOP* holds between two neighbouring small open NMS of the EU, engaged in wheat trade. All cointegration tests showed a strong long-run relationship between Hungarian and Slovenian wheat producer prices. More importantly, however, all three methods employed rejected the *LOP* hypothesis (i.e., rejected H_1), and thus the unit elasticity. This has important international business implications and is in line with most empirical studies, which have failed to show the validity of the *LOP*, even in a single country/multi-regions setting (e.g. Brosig et al. 2011, Eryigit & Karaman, 2010). Listorti and Esposti (2012) argue that since the concept of *LOP* is a static one, with quite restrictive assumptions, it is unlikely to be observable in practice for longer periods. Price causality was tested in the linear VEC model and Gonzalo-Pitarakis VEC model alike, both strongly concluding that Hungarian wheat producer prices are determining Slovenian ones (i.e., not rejecting the H_2 null). A rather rarely employed Enders and Siklos (2001) cointegration test, which is capable of detecting short-run adjustment asymmetries, could not reject the symmetric short-run price adjustment as response for positive or negative deviations from long-run equilibrium, thus suggesting competitive market conditions (i.e., not rejecting H_3 null). We also do not reject H_4 null, i.e. that price elasticities are trade volume dependent, albeit the difference between the regime dependent coefficients is rather small.

Therefore, despite the similar evolution of Hungarian and Slovenian wheat producer prices, the existing long-run cointegration, common EU membership and non-the-less being neighbouring countries with rather favourable transport infrastructure (railway and motorway connections) the *LOP* does not hold. At first glance, it might seem to be a surprising result, considering that the Slovenian port Koper is also one of the important gateways of the Hungarian wheat export, and the very large share Hungary possess in total Slovenian wheat imports. Manrai and Manrai (2001), based on the demographic-economic demand potential and the availability of logistics infrastructure as attractiveness for international marketing operations, efficient distribution and promotion of products, they separated Hungary and Slovenia in two different clusters: Hungary in Cluster 1 (the most promising group of countries attractive on both dimensions) and Slovenia in Cluster 2 (either moderate in terms of attractiveness on both dimensions or countries with a trade-off between two dimensions; one high, other low). This difference can be one of the reasons that the *LOP* does not hold.

Considering the competitiveness of the European wheat market, however, we may argue that:

- regardless of the extreme importance it plays for Slovenia in this respect, Hungary is just one of many suppliers on the international wheat market, which is dominated by much larger players (e.g. Germany, Russia, Ukraine and some overseas countries).
- Previous research emphasized that Hungarian wheat producer prices are caused by the largest buyer, i.e., Germany (Bakucs et al., 2012). It is plausible to accept, that despite causality tests persistently emphasizing the information running from Hungary towards the neighbouring Slovenia, wheat producer prices in both countries are actually determined by the regional (e.g. German and possibly some other regional) and global wheat markets, through a 'hidden' causality relationship, not the subject of this paper.
- Except German, Austrian, Italian and perhaps Dutch destinations, Hungarian wheat exports run through the Slovenian port of Koper, thus it is reliant on wheat international trade running through the Slovenian port.

- Despite its considerable role in the Slovenian wheat imports, and the relative importance on the European wheat market, Hungarian wheat producer market remains typical for a small open economy, with a price taking behaviour on the regional and other global wheat markets. Thus, despite its very large share in Slovenian wheat imports it cannot exercise market power, or alter prices set on the international market.

7. CONCLUSIONS

In this paper, we have analyzed the dynamics of spatial wheat producer price transmission between neighbouring Hungary and Slovenia by testing for the *LOP*, identifying the price-leading market, testing for adjustment asymmetries and finally, the role trade quantities may play in wheat producer price market integration. The analysis covers the January 2000-April 2011 period, characterized by rapidly changing market conditions due to the adjustment and membership of Hungary and Slovenia to the 2004 enlarged EU and the Slovenian adoption of the Euro on January 1st, 2007. In addition, during the years 2009-2011 both Hungary and Slovenia have faced economic recession, macroeconomic and sectorial adjustments, including in wheat producer markets. Despite these dramatic changes in the political, social and macroeconomic environment, long-run wheat producer price relationships and wheat producer price market integration do not seem to have been affected. As a result, a linear and non-linear VEC model were found to be capable of adequately depicting long-run wheat producer price relationship between Hungary and Slovenia.

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References

Bakucs, L. Z., Brümmer, B., von Cramon-Taubadel, S., & Fertő, I. (2012). Wheat market integration between Hungary and Germany. Applied Economics Letters, *19*(8), 785–788.

Balcombe, K., Bailey, A., & Brooks. J. (2007). Threshold effects in price transmission: the case of Brazilian wheat, maize and soya prices. *American Journal of Agricultural Economics*, *89*(2), 308–323.

Bessler, D. A., Yang, J., & Wongcharupan, M. (2003). Price dynamics in the international wheat market: modelling with error correction and directed acyclic graphs. *Journal of Regional Science*, *43*(1), 1–33.

Brosig, S., Glauben, T., Götz, L., Weitzel, E. B., & Bayaner, A. (2011). The Turkish wheat market: spatial price transmission and the impact of transaction costs. *Agribusiness*, *27*(2), 147–161.

Bukenya, J. O., & Labys, W. C. (2005). Price convergence on world commodity markets: fact or fiction? *International Regional Science Review*, *28*(3), 302–329.

Carew, R.. & Florkowski, W. J. (2003). Pricing to market behaviour by Canadian and U.S. agri-food exporters: evidence from wheat, pulse and apples. *Canadian Journal of Agricultural Economics*, *51*(2), 139–159

Dawson, P. J., Sanjuan, A. I., & White, B. (2006). Structural breaks and the relationships between barley and wheat futures prices on the London international future exchange. *Review Agricultural Economics*, *28*(1), 585–594.

Ejrnæs, M., & Persson, P. G. (2000). Market integration and transport costs in France 1825–1903: a threshold error correction approach to the law of one price. *Extrapolation in Economic History*, *37*(2), 149–173.

Enders, W., & Siklos, P. L. (2001). Cointegration and threshold adjustment. *Journal of Business & Economic Statistics*, *19*(2), 166–176.

Engle, R. F., & Granger, C. W. J. (1987). Co-integration and error correction: representation, estimation and testing. *Econometrica*, *55*(2), 251–276.

Eryigit, K. Y., & Karaman, S. (2011). Testing for spatial market integration and law of one price in Turkish wheat markets, *Quality & Quantity*, *45*, 1519–1530.

Eurostat. (2012). Comext trade database. Luxembourg: Eurostat. Available at: http://epp.eurostat.ec.europa.eu/newxtweb/.FAO. (2012). FAOSTAT commodities production and trade database. Rome: Food and Agriculture Organization of the United Nations. Available at http://www.fao.org/economic/ess/ess-trade/ar/.

Ghoshray, A. (2002). Asymmetric price adjustment and the world wheat market. *Journal of Agricultural Economics*, *53*(2), 299–317.

Ghoshray, A. (2007). An examination of the relationship between U.S. and Canadian durum wheat prices. *Canadian Journal of Agricultural Economics*, *55*(1), 49–62.

Ghoshray, A. (2010). Smooth transition effects in price transmission: the case of international wheat export prices. *Economics Letters*, *106*(3), 169–171.

Gonzalo, J., & Pitarakis, J. (2006). Threshold effects in cointegrating relationships. *Oxford Bulletin of Economics and Statistics*, *68*(Supplement), 813–833.

Götz, L., & von Cramon-Taubadel, S. (2008). Considering threshold effects in the long-run equilibrium in a vector error correction model: an application to the German apple market. Paper presented at the 12th Congress of the European Association of Agricultural Economists, August, 26-29, 2008, Ghent, Belgium.

Granger, C. (1981). Some properties of time series data and their use in econometric model specification. *Journal of Econometrics*, *16*(1), 121–130.

Jin, H. J., & Miljkovic, D. (2008). Competitive structure of U.S. grain exporters in the world market: a dynamic panel approach. *Journal of International Economic Studies*, *12*(1), 33–63.

Johansen, S. (1988). Statistical analysis of cointegrating vectors. *Journal of Economic Dynamics and Control*, *12*(2-3), 231–254.

Johansen, S., Mosconi, R., & Nielsen, B. (2000). Cointegration analysis in the presence of structural breaks in the deterministic trend. *Econometrics Journal*, *3*(2), 216–249.

Listorti, G., & Esposti, S. (2012). Horizontal price transmission in agricultural markets: fundamental concepts and open empirical issues. *Bio-based and Applied Economics*, *1*(1), 81–96.

Manrai, L. A., & Manrai, A. K. (2001). Marketing opportunities and challenges in emerging markets in the new millennium: a conceptual framework and analysis. *International Business Review*, *10*(5), 493–504.

Mainardi, S. (2001). Limited arbitrage in international wheat markets: threshold and smooth transition cointegration. *Australian Journal of Agricultural and Resource Economics*, *45*(3), 335–360.

Mohanty, S., & Langley, S. (2003). The effects of various policy regimes in the integration of North American grain markets. *Canadian Journal of Agricultural Economics*, *51*(1), 109–120.

Pall, Z., Perekhozhuk, O., Teuber, R., & Glauben, T. (2013). Are Russian wheat exporters able to price discriminate? Empirical evidence from the last decade. *Journal of Agricultural Economics*, *64*(1), 177–196.Pick, D. H., & Park, T. A. (1991). The competitive structure of U.S. agricultural exports. *American Journal of Agricultural Economics*, *73*(1), 133–141.

Pick, D. H., & Carter, C. A. (1994). Pricing to market with transactions denominated in a common currency. *American Journal of Agricultural Economics*, *76*(1), 55–60.

Saikkonen, P., & Lütkepohl, H. (2000). Testing for the cointegrating rank of a VAR process with an intercept. *Econometric Theory*, *16*(3), 373–406.

SORS. (2011). SI-STAT Data Portal. Ljubljana: Statistical Office of the Republic of Slovenia. Available at http://pxweb.stat.si/pxweb/Dialog/statfile1.asp.

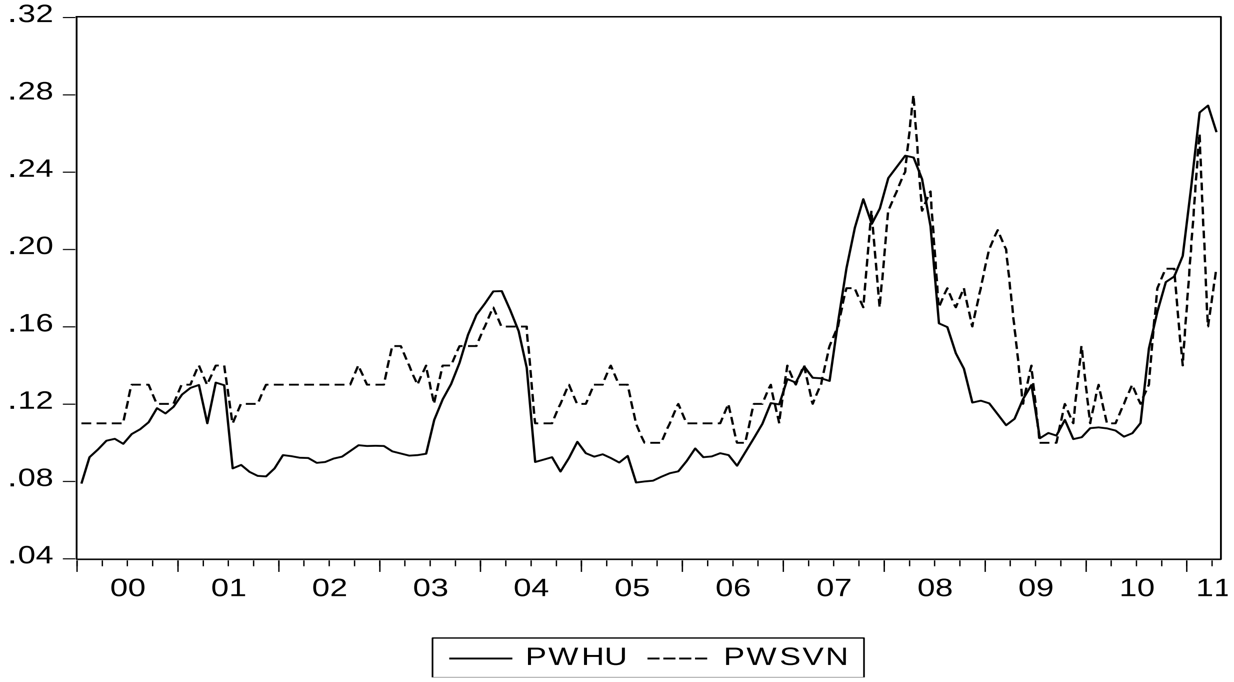
Stephens, E., Mabaya, E., von Cramon-Taubadel, S., & Barrett, C. B. (2008). Spatial price adjustment with and without trade. Paper presented at American Agricultural Economics Association Annual Meeting, July, 27-29, 2008, Orlando, Florida.

Tun-Hsiang, E. Y., Bessler, D. A., & Fuller, S. W. (2007). Price dynamics in U.S. grain and freight markets. *Canadian Journal of Agricultural Economics*, *55*(3), 381–397.

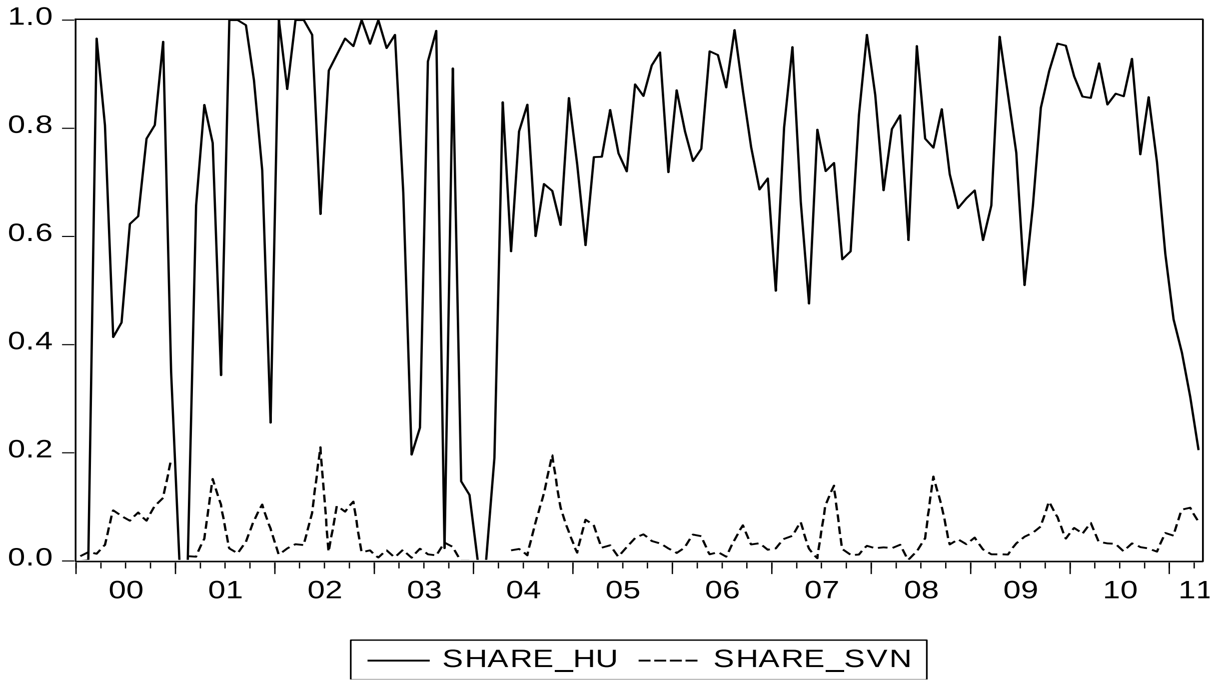
Thompson. S. R., Sul, D., & Bohl, M. T. (2002). Spatial market efficiency and policy regime change: seemingly unrelated error correction estimation. *American Journal of Agricultural Economics*, *84*(4), 1042–1053.

Tomek, W., & Robinson, K. J. (2003): Agricultural product prices. Ithaca: Cornell University.

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), 1–18.



Source: Hungarian and Slovenian Statistical Office data.
Figure 1 Hungarian (PWHU) and Slovenian (PWSVN) wheat producer prices (in Euro per kg)



Note: SHARE_HU – share of Hungarian wheat in total Slovenian wheat imports
SHARE_SVN – share of wheat exports to Slovenia in total Hungarian exports
Figure 2 Importance of wheat trade between Slovenia and Hungary

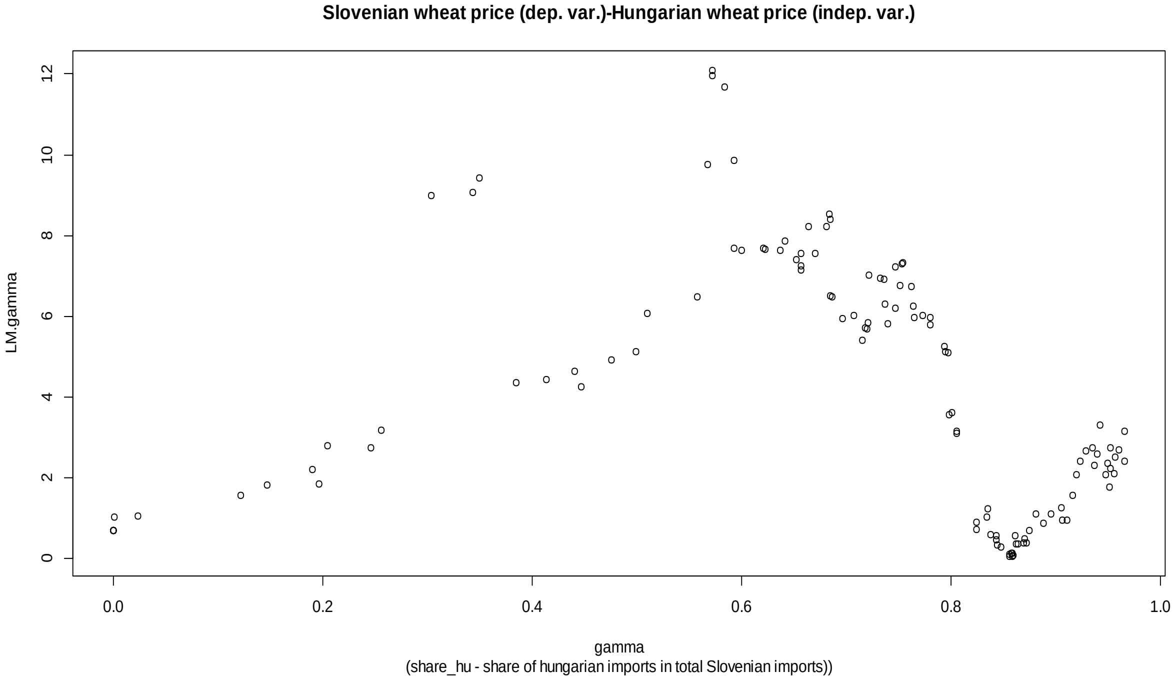


Figure 3 Gonzalo-Pitarakis sup LM_γ Statistics with respect to the share of Hungarian imports in total imports into Slovenia

TABLE 1. Descriptive Statistics, 2000-2011

	PWHU*	PWSVN*	SHARE_HU#	SHARE_SVN#
Mean	0.125915	0.141029	0.713118	0.046633
Median	0.107472	0.130000	0.780450	0.031243
Maximum	0.274447	0.280000	1.000000	0.210028
Minimum	0.078847	0.100000	0.000000	0.000000
Std. Dev.	0.046771	0.034881	0.262306	0.041449
Skewness	1.517905	1.512270	-1.317965	1.606474
Kurtosis	4.459553	5.309191	4.037144	5.670121
Jarque-Bera	64.29651	82.05446	45.46814	95.26202
Probability	0.000000	0.000000	0.000000	0.000000
Observations	136	136	136	131

Note: * expressed in Euro, PWHU – Hungarian wheat producer prices, PWSVN – Slovenian wheat producer prices.
expressed as %, SHARE_HU – the share of Hungarian wheat imports in total Slovenian imports, SHARE_SVN – the share of Slovenian wheat exports in total Hungarian wheat exports. Source: Own estimations based on Hungarian and Slovenian Statistical Office data, and Eurostat Comtrade database

TABLE 2. Augmented Dickey-Fuller unit root test results

Variable	Deterministic specification	
	Constant	Constant and trend
PWHU	-1.527	-2.068
PWSVN	-2.588*	-2.813
SHARE_HU	-7.193***	-7.119***
ΔPWHU	-9.174***	-
ΔPWSVN	-16.174***	-
ΔSHARE_HU	- 15.000***	-

Note: lag length was determined by the Schwarz-Bayesian Criteria, * significant at 10%, ** significant at 5%, ***significant at 1%.
Source: Own estimations based on Hungarian and Slovenian Statistical Office data.

TABLE 3. Linear cointegration test results

Number of cointegration (CI) vectors	p–value
Johansen trace test	
0	0.058
1	0.601
Saikkonen and Lütkepohl test	
0	0.007
1	0.821

Note: lag length was chosen by the Schwarz-Bayesian information criteria, one constant was included in the cointegration space.
Source: Own estimations based on Hungarian and Slovenian Statistical Office data.

TABLE 4. Linear Vector Error Correction model

Cointegrating Eq:	Coint Eq (1)	
PWSVN $_{t-1}$	1.000000	
PWHU $_{t-1}$	-0.558390 [-6.13949]	
constant	0.792005 [4.03416]	
Error Correction:	**Δ (PWSVN)**	**Δ (PWHU)**
α	-0.358996 [-4.10993]	-0.064199 [-0.96248]
Δ (PWSVN) $_{t-1}$	-0.264402 [-2.76906]	-0.062590 [-0.85839]
Δ (PWHU) $_{t-1}$	0.224485 [1.74773]	0.263592 [2.68742]
R^2	0.268217	0.073041
Adj. R^2	0.257044	0.058889
Sum sq. resids	1.761827	1.027376
S.E. equation	0.115970	0.088558
F-statistic	24.00737	5.161186
Log likelihood	100.0720	136.2080
Akaike IC	-1.448836	-1.988178
Schwarz IC	-1.383958	-1.923301
Log likelihood		248.1146
Akaike IC		-3.568875
Schwarz IC		-3.374244

Note: Student t-statistics in parentheses.
Source: Own estimations based on Hungarian and Slovenian Statistical Office data.

TABLE 5. Enders and Siklos (2001) non-linear cointegration tests

Variable	Coefficient	Std. Error
Above Threshold (ρ_1)	-0.282521	0.099458
Below Threshold (ρ_2)	-0.374776	0.124350
Differenced Residuals(t-1)*	-0.227293	0.101566
Differenced Residuals(t-2)*	-0.005660	0.093698
Threshold value (Tau):	0.000000	
F-equal:	0.406430	(2.460716)#
T-max value:	-2.840611	(-2.171301)#
F-joint (Φ):	7.280372	(5.873208)#

* Information criteria (AIC, SBC, HQ, FPE) have unanimously selected 2 lags

\# Simulated 5% critical values (10,000 replications)

Source: Own estimations based on Hungarian and Slovenian Statistical Office data.

TABLE 6. Threshold Vector Error Correction Model

Cointegrating Eq:	Coint Eq (1) (SHARE_HU<57.2%) 25 obs.		Coint Eq (2) (SHARE_HU>=57.2%) 108 obs.	
$PWSVN_{t-1}$	1.000		1.000	
$PWHU_{t-1}$	-0.497		-0.650	
	[-3.714]		[-6.696]	
constant	1.015		0.572	
	[3.387]		[2.687]	
Error Correction:	Δ (PWSVN)	Δ (PWHU)	Δ (PWSVN)	Δ (PWHU)
α	-0.574284	0.004417	-0.370825	-0.067145
	[-1.93141]	[0.02441]	[-4.09765]	[-0.86319]
$\Delta (PWSVN)_{t-1}$	-0.528543	-0.078526	-0.168906	-0.020986
	[-2.04267]	[-0.49855]	[-1.51770]	[-0.21938]
$\Delta (PWSVN)_{t-2}$	-0.473408	0.025464	0.206418	0.159653
	[-1.95054]	[0.17236]	[1.92106]	[1.72860]
$\Delta (PWHU)_{t-1}$	0.825582	0.466203	0.23933	0.2228
	[1.84115]	[1.70798]	[1.83845]	[1.99110]
$\Delta (PWHU)_{t-2}$	-0.186394	0.047955	-0.0988	-0.081887
	[-0.55133]	[0.23302]	[-0.69912]	[-0.67411]
R^2	0.531283	0.064777	0.288648	0.087232
Adj. R^2	0.43754	-0.122268	0.261023	0.051784
Sum sq. resids	0.358243	0.132745	1.168901	0.863635
S.E. equation	0.133836	0.081469	0.10653	0.091569
F-statistic	5.667424	0.346318	10.4487	2.460882
Log likelihood	17.59429	30.00401	91.16228	107.5064
Akaike IC	-1.007543	-2.000321	-1.595598	-1.898266
Schwarz IC	-0.763768	-1.756546	-1.471425	-1.774094
Log likelihood		60.80192		206.8094
Akaike IC		-3.824154		-3.589062
Schwarz IC		-3.190338		-3.266213

Note: Student t-statistics in parentheses.
Source: Own estimations based on Hungarian and Slovenian Statistical Office and Eurostat data.

TABLE 7. Regime dependent descriptive statistics of the SHARE_HU variable

Statistic	Smaller regime ‘extreme’	Larger regime ‘normal’
Mean	0.256805	0.820974
Median	0.250691	0.840369
Maximum	0.567660	1.000000
Minimum	0.000000	0.571991
Std. Dev.	0.197725	0.121830
Relative Std. Dev.	0.769942	0.148396
Observations	26	110

Source: Own estimations based on Hungarian and Slovenian Statistical Office data.