Wheat market integration between Hungary and Germany

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Abstract

One of the most important targets of the Common Agricultural Policy (CAP) is to facilitate the spatial integration of agricultural markets within the individual member states as well as within the Community. On a spatially integrated market, price information should freely flow between member states. According to the European Commission, national Governments and their regulations should help to attain the goal of a common, integrated, and efficient market. For a small open economy, such as Hungary, market efficiency, and market information flow has at least two important political consequences. The first one is the transmission of prices by some actors of the chain either vertically or spatially. This issue is quite relevant for Hungary, considering the structure of its agri-food market. The second problem relates to the national agricultural support system completing the CAP in the New Member States (NMS). This paper focuses on the first topic, by testing for price transmission between German and Hungarian producer prices. Given the changing nature of market conditions over the past five years, a flexible Markov-Switching model for price transmission is proposed and estimated for the analysis of price transmission between Hungarian and German wheat.

Keywords: wheat market integration, markov switching modell

JEL kód: Q11

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1. Introduction

Hungary is the major wheat exporting country in the group of the ten countries in the 2004 enlargement. The EU-15 countries, in particular Germany and Austria, are traditionally important trading partners. The wheat trading relations became more important after the accession in 2004, when the excellent harvest led to a quick filling of the available intervention storage capacities, and hence trade into other countries' intervention took place, with strong impact on regional price levels. This is but one example of the many effects of European cereal policy on the price transmission in spatially separate markets. The resulting price levels are important for the agricultural sector because cereal prices, in particular wheat prices, are important on both the revenue (wheat sales) and the cost (wheat for feeding) side of many farms' balance sheets. On the demand side, wheat is important for industrial and bioenergy uses, besides its use for concentrates, and human nutrition. Hence, the development of the price transmission for wheat between Hungary and Germany is an important topic for market analysis.

There exists some literature on spatial price transmission in wheat markets, mainly focusing on US-Canada relationships, and international markets (e.g. Bessler et al. 2003; Ghoshray 2002 and 2007; Tun-Hsiang et al. 2007; Mainardi 2001; Mohanty and Langley 2003). However, there are just a few papers on the European wheat markets (Dawson et al. 2006; Ejrnæs andPersson 2000; Thompson et al. 2002). To the best of our knowledge, there has been no published research focusing on spatial integration of cereal prices between a CEE country and EU 15. Because of the deficits in the development of market institutions, and market inefficiencies, the evolution of spatial price transmission is perhaps of even more interest in transition than in developed economies. On the other hand, the higher variability in market conditions implied by the lack or underdevelopment of market institutions might lead to multiple price relations over time: The law of one price (net of trade costs), which should be valid in the absence of market frictions, is unlikely to be empirically observable at every single point in time.

This paper adds to the existing literature by analysing the pattern of price transmission between Hungarian and German wheat at the producer level, allowing for the combination of multiple price relationships in the framework of a switching regime model. We utilize weekly wheat price data from January 2003 to September 2007 in order to estimate Vector Error Correction and Markov-Switching Error Correction methods. The paper is organised as follows. Section 2 briefly describes the methodology, section 3 presents the results of the empirical analysis. Section 5 links the results to an analysis of trade flows, before section 5 provides a summary.

2. Methodology

The cointegration framework is appropriate when using non-stationary time series. Most commonly used are linear cointegration tests, followed by the estimation of a Vector Error Correction Model, VECM. Johansen et al. (2000) generalised the Johansen (1988) maximum likelihood cointegration test in order to include up to two breaks. The procedure estimates the following model:

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

where Y_t is a vector of non-stationary variables (in our case the German and Hungarian wheat 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 speed of adjustment parameter matrix, β are the long run cointegration coefficients and μ are the long run drift parameters. 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 can be tested using likelihood ratio tests.

A more flexible, yet slightly difficult approach is to allow the price equation system parameters to vary according to the possible shifts in the data generating process. Threshold models allow defining two or more regimes with regime dependent short-run parameters and adjustment coefficients. Threshold models are often used to test price integration, since the threshold may be interpreted as transaction costs. Hamilton (1989) developed the Markov-switching vector autoregressive model. The advantage of Markov-switching (MS) class models is that it allows time series analysis with different regimes, when the corresponding state variable is not known. In this paper we apply Markov-switching error correction models, MSVECM, allowing shifts in the short-run parameters, intercept, and residual variance according to the state of the system:

$$\Delta Y_t = v(s_t) + \alpha(s_t)(\beta^{'} Y_{t-1}) + \sum_{i=1}^{k} D_i(s_t)\Delta Y_{t-i} + u_t \qquad (2)$$

where Y_t is the non-stationary price vector, v is the vector of intercept terms, α is the vector of the speed of adjustment coefficients, and β is the long-run cointegrating vector. D_i are the autoregressive, (short-run parameters) matrices. As before, u_t are assumed to have the usual properties. s_t is the state variable, where $s_t = 1,...,M$ indicates in which of the M possible regimes the system might be in. The state of the system, however, is not directly observed. Generally, the probability of the system of being in state s_t might depend on the full history of the system. In MS modelling, the following simplifying assumption is made:

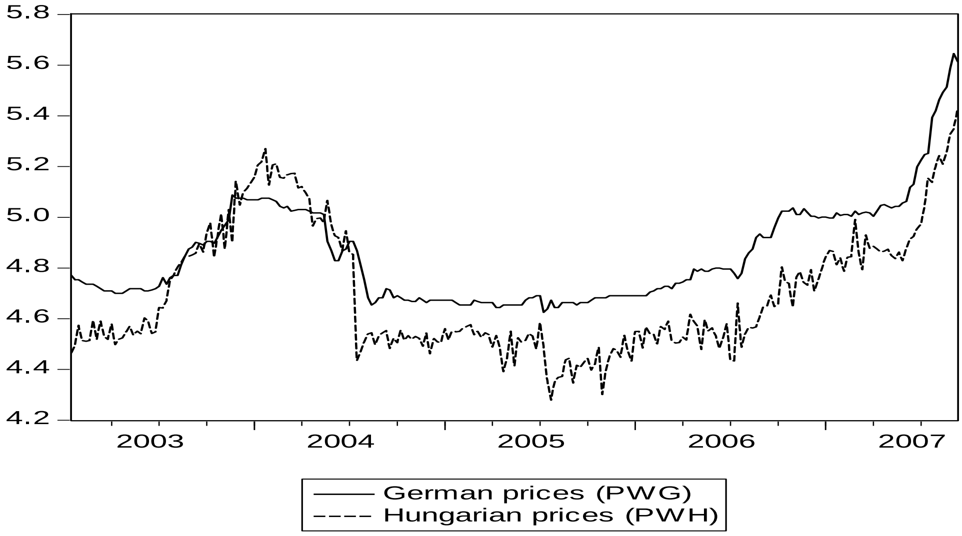
$$\Pr(s_t \mid S_{t-i}, \Delta Y_{t-1}, \beta^{'} Y_{t-1}) = \Pr(s_t \mid s_{t-1,} \Pi)$$

where Π is the matrix of transition probabilities, i.e., the probability of today's state does functionally depend only on the state in the previous time period. Estimation of MS time series models is usually by variants of an Expectation-Maximization (EM) algorithm, e.g., available in the MSVAR package of Krolzig (2004), for the Ox programming language.

3. Empirical analysis

The data (in logs) for the empirical analysis is presented in Figure 1. Weakly German (PWG) and Hungarian (PWH) prices, between January 2003 and September 2007, totalling 243 observations were used. Data was provided by the Agricultural Economics Research Institute (AKII).

Figure 1. German and Hungarian wheat producer prices based on data from AKII



Vector Error Correction Model, VECM

Since the time series for wheat prices were found to be non-stationary[1], the cointegration framework is generally suitable for further analysis. Linear cointegration (CI, Johansen, 1988) tests could not reject the null hypothesis of no CI. The framework of Johansen et al. (2000) presented in section 2, allowing the inclusion of up to 2 structural breaks in the long-run relationship, provides a more flexible extension. The breakpoints should be known *a priori*, the test is not capable of endogenously searching for structural breaks[2]. An obvious choice for the time of the break points would be the date of the level shifts in the individual series, identified through the Perron (1997) unit root test. Table 1 presents the results of the Johansen (2000) CI test, using observation 79 as the break point.

[1] Unit root test results are available from authors upon request. Unit root tests in the presence of structural breaks revealed a break in the Hungarian price occurring in July 2004.
[2] We also applied Gregory and Hansen (1996) methodology which endogenously searches for possible structural breaks, but inconclusive results obtained.

Table 1. Johansen (2000) CI tests using t=79 as break point[3]

No. of CI vectors	Trace statistic	Significance
0	25.91	0.035
1	5.39	0.552

The null of no cointegration was rejected in favour of the alternative hypothesis of cointegration with a structural break occurring at observation 79, July 2004. Normalised on German wheat prices, the long-run relationship is (*t* statistics in brackets):

PWG = 1.007 + 0.108D +0.797PWH,

(- 2.92) (- 2.95) (- 10.86)

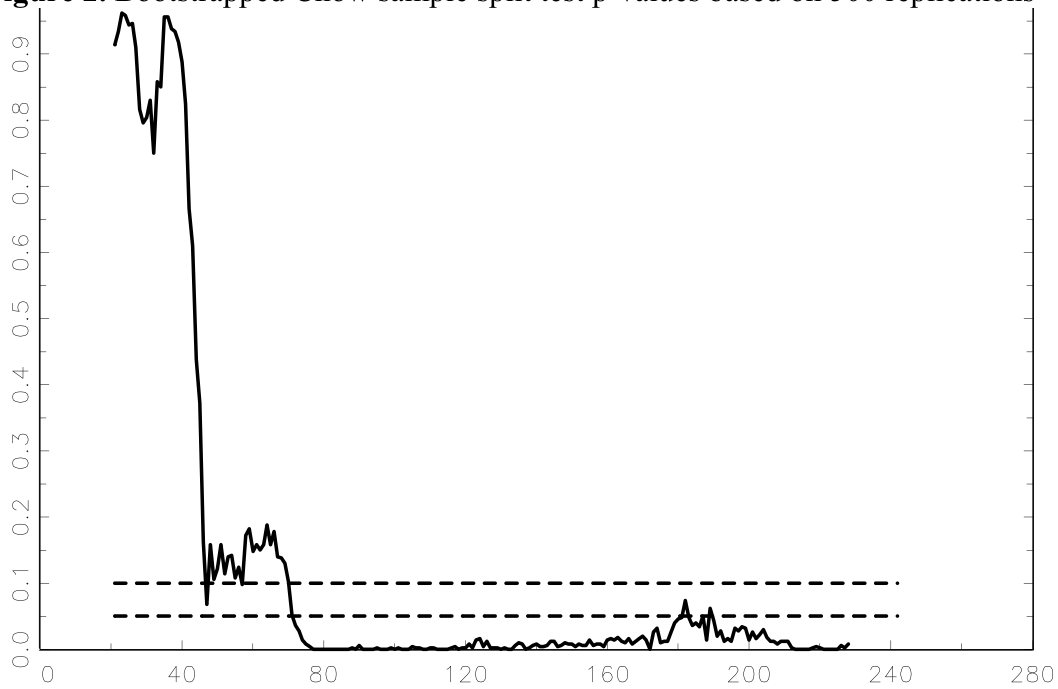
where $D = \begin{cases} 1 & if \quad t > 79 \\ 0 & otherwise \end{cases}$.

The date of the structural break in July 2004, coincides with the start of the harvest in Hungary. The 2004 harvest was exceptionally good, not only in Hungary, but also worldwide. The news of the good harvest, combined with the lack of storage facilities drove prices down, causing the level shift. The adjustment coefficients (α) for the system of equations are (t statistics in brackets): α_{PWG}=0.013 (0.64), and α_{PWH}=0.261 (4.26). Since the long-run relationship was normalised on German prices, the adjustment coefficient of the Hungarian prices has the correct sign, and it is also significant. The adjustment coefficient of German prices is not significantly different from 0. It follows that, as expected, Hungarian prices do adjust to German prices, and not other way around. The residuals of the estimated VECM model do not seem to suffer from autocorrelation up to lag 42 with $\chi^2(42)$=52,87 (p=0.121). However, the distribution of residuals is likely non-normal (Jarque-Bera test with p=0.00). The null hypothesis of the law of one price, i.e. equality of the coefficients of German and Hungarian prices in the long-run relationship is also rejected. Even though there exist a long-run linear relationship between prices, the system might not be stable. Chow tests are used to check for system stability. Since the small sample distribu-

[3] Helmut Lütkepohl's JMulti software was used to run the Johansen (2000) cointegration tests

tions of the test statistic under the null hypothesis may be different from the asymptotic χ^2 distributions (Candelon and Lütkepohl, 2000), bootstrapped p-values are used. Figure 2 presents the bootstrapped p-values of the Chow sample split test, where each observation was considered as a possible break date.

Figure 2. Bootstrapped Chow sample split test p-values based on 500 replications



Most of the p-values are below the 5% critical level, strongly suggesting the instability of the system. Hence, the linear model is not appropriate and a more flexible representation should be used instead.

Markov switching vector error correction model, MS-VECM

Several specifications of MS-VECM were considered, including a restricted one with given cointegration vector (the long-run relationship identified in the previous section), however a completely flexible, unrestricted model with regime-dependent long-run relation finally was preferred on grounds of Akaike information criterion, and inspection of the residuals. The estimated MSIAH(3)-AR(3) allows for shifts in the intercept, mean, autoregressive parameters and residuals across regimes. AIC and log-likelihood criteria were used to determine the lag length and the number of regimes: Three lags in first differences, and three regimes were selected. A formal likelihood ratio test of the null hypothesis of linearity against the alternative of non-linear representation rejected the linearity null (p=0.00). Generally, the MSVECM(3)-AR(3) model appears to be well specified (see figures 5 and 6); there is no evidence for autocorrelation, $\chi^2(49)=57.49$ (p=0.18), 52.03 (p=0.35) and 49.78 (p=0.44) in the 3 regimes respectively. Homoskedasticity of the residuals cannot be rejected, $\chi^2(18)=12.77$ (p=0.805), 21.39 (p=0.259), and 16.15 (p=0.581) for the 3 regimes respectively. However, normality is rejected for all regimes. Table 2 presents the characteristics of the identified regimes.

Table 2. Regime properties

Regimes	Indicative labelling	No. of obs.	Prob.	Duration
Regime 1	*"Great uncertainty"*	13.6	0.05	1.27
Regime 2	*"Law of one price"*	65.6	0.27	6.96
Regime 3	*"Normal"*	159.8	0.67	17.07

Regime 3 contains most observations, and also has the longest duration and highest probability, therefore we call it 'Normal regime'. Regime 2 has a shorter duration, containing 27% of observations, with an average duration of 7 weeks. The label 'Law of one price regime' arises from the estimated long-run parameters (see below). Finally, regime 3 is the least stable with the shortest duration (less than 2 weeks on average), and contains only 13 observations, we call it the regime of "Great Uncertainty'.

Table 3. Matrix of transition probabilities

Regimes	Regime 1	Regime 2	Regime 3
Regime 1	0.212	0.274	0.513
Regime 2	0.105	0.856	0.038
Regime 3	0.023	0.035	0.941

The matrix of transition probabilities, in table 3, presents the probabilities of transition of the system from one regime into another. Figures on the diagonal represent the probabilities of the system remaining in the actual regime. The more stable regimes (2 and 3), have high probabilities in their respective column, whilst regime 1, the most unstable, shows only low probabilities:

At any point in time, regime 1 is an unlikely alternative for the next period. If a regime change happens, from regime one the system will most likely shift to regime 3 (51% probability), the probability of the system moving to regime 2, being substantially lower (27%). From the more stable regime 2, if a change occurs, the system will most likely move to the *Great uncertainty regime* (but only with 10% probability). The system is stable in the *Normal regime*, but if a change happens, it will move most likely – albeit with low probability – to regime 2 (4%).

Table 4. MS vector error correction model, dependent variable ΔPWH

	Regime 1		Regime 2		Regime 3	
	Coeffi-cient	t - stat	Coefficient	t - stat	Coefficient	t - stat
Constant	- 4.268	-585.41	- 0.014	- 0.158	- 0.201	- 1.688
ΔPWH_{t-1}	- 0.207	- 43.148	- 0.453	- 5.992	- 0.634	- 7.915
ΔPWH_{t-2}	0.442	52.089	- 0.056	- 1.191	- 0.471	- 5.58
ΔPWH_{t-3}	- 1.151	- 117.72	- 0.043	1.075	- 0.082	- 1.126
ΔPWG	3.117	236.01	- 0.833	- 6.24	0.366	1.728
ΔPWG_{t-1}	- 0.092	- 10.417	0.311	1.78	0.343	1.459
ΔPWG_{t-2}	- 1.518	- 98.552	0.368	2.049	- 0.009	- 0.043
ΔPWG_{t-3}	- 4.366	- 468.9	- 0.368	- 2.333	0.458	1.944
	vector of long-run parameters normalised on Hungarian prices					
Intercept	5.597	-585.41	0.091	- 0.158	4.025	- 1.688
PWG	- 2.104	434.07	- 0.988	4.474	- 1.861	1.86
Res. std. errors	0.00114		0.01964		0.04827	

Because of the very low number of observations, coefficient estimates (table 4) in the first regime should be treated with caution. This is also shown by the unusually low standard errors and high t statistics. The alternative of using only two regimes was not supported by a comparison of the AIC criteria, and also the first regime seems to capture the uncertainties while there is disarray in the price relationship. Using the results from table 4, the Hungarian – German price relationship may be characterised as follows:

1. The *great uncertainty regime* exhibits very low residual standard errors (mostly due to the low number of observations in this regime) in combination with large price changes. The adjustment process is very fast, 76% of the price difference is corrected in a week. However, the long-run relationship gives indicate a price elasticity of -2.1, i.e., a percentage change in German

wheat prices causes a relative change in Hungary of about the double magnitude. The upper panel of figure 3 shows the first difference of the Hungarian prices. If compared to the probabilities of the system being in the *great uncertainties regime* (second panel), it can be seen that the price relationship moves into this regime when significant negative price differences arise.

2. The *law of one price regime* is characterized by the validity of the law of one price between Germany and Hungary in the long-run. The corresponding price coefficients are equal, there is no significant relative margin between German and Hungarian prices, hence, there is a perfect price information flow. 16% of a deviation from the equality of prices, i.e., the long-run relationship in this regime, is adjusted within the following period. Most of the short-run coefficients are statistically significant. Comparing the first and the third panel in Figure 3 reveals that this regime coincides mostly with periods of small changes in Hungarian prices.

3. *Normal regime*: the law of one price does not hold, there is a large absolute constant margin between prices, possible to interpret as transaction costs. Short-run coefficients are not as significant as in regime 2. The adjustment is slow, only 5% of the price difference is adjusted during a week.

Figure 3. Hungarian price differences and regime probabilities

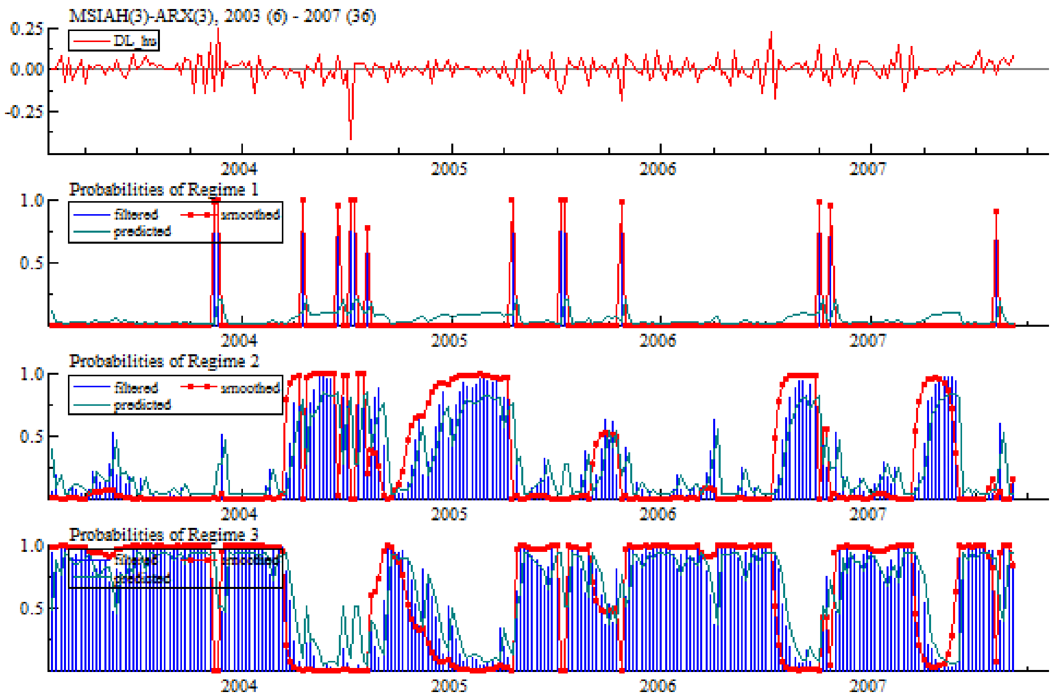
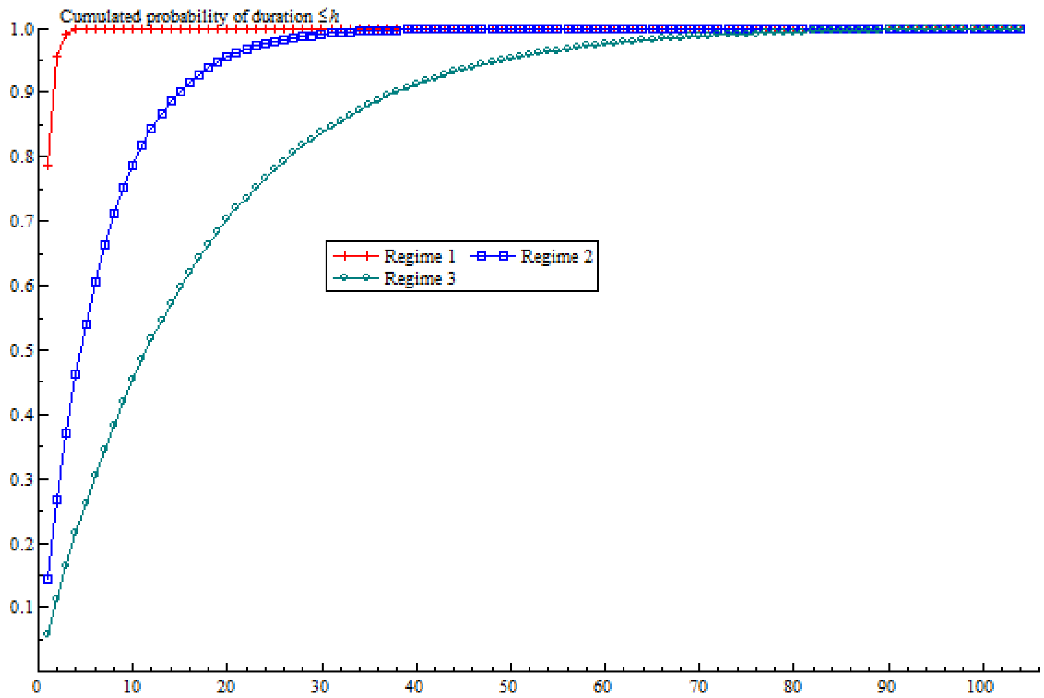


Figure 4 show the cumulative probabilities for the duration of regime 1, 2 and 3. It can be seen that the duration of regime 1 is less than 2 weeks, whilst the duration of regime 2 is substantially longer. The most stable regime is regime 3, the probability of observing for more than 10 weeks is 50%.

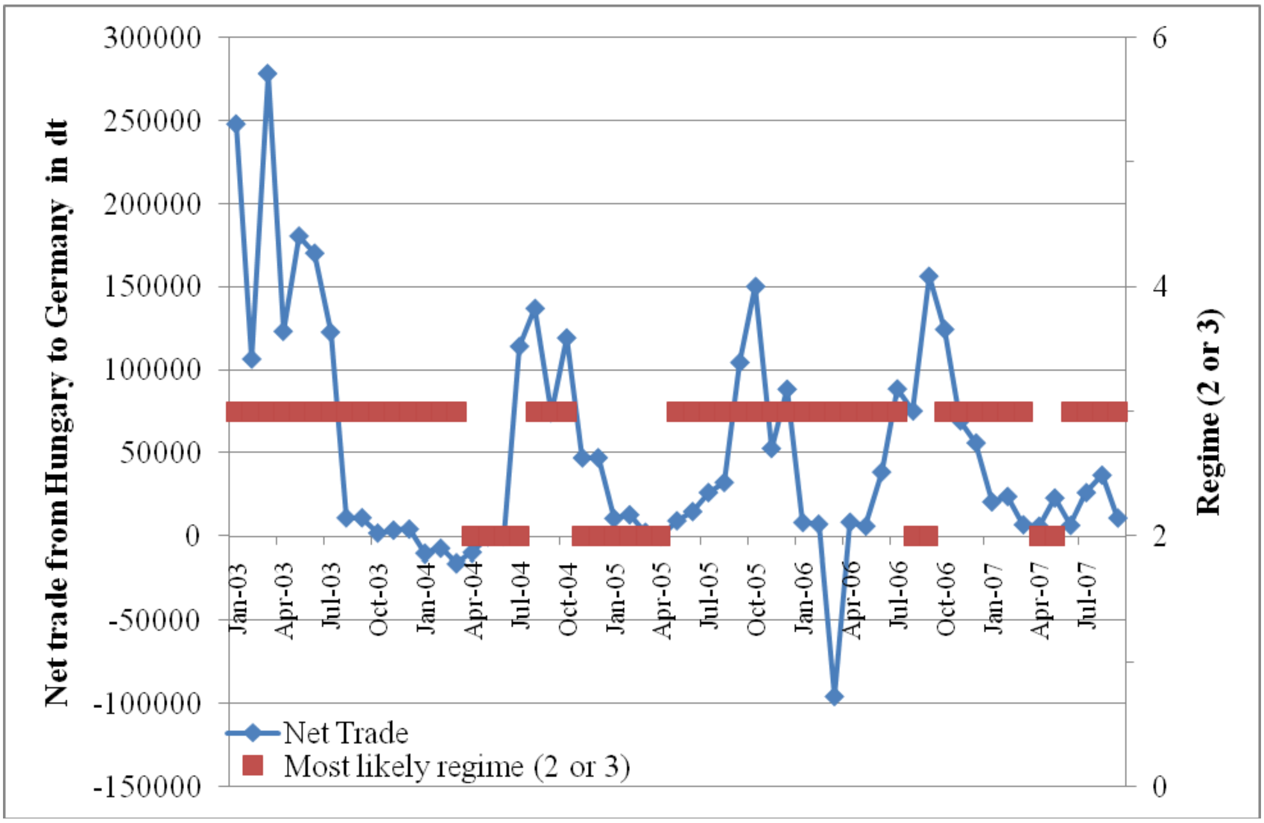
Figure 4. Cumulative probabilities for duration of regimes less than *h* weeks



4. Trade analysis

The regime classification according to the above results (figure 3) exhibits sufficient discriminatory power in the sense that for most periods, one regime probability is close to one while the remaining ones are close to zero. The interesting point is, however, can this classification be somehow linked to the degree of market integration, e.g., to the volume of trade, or to specific events over time? In particular, the volume of trade should exhibit some kind of relationship to the estimated regime relationships if the latter are more than an statistical artefact. Bilateral trade data were obtained from Eurostat's Comext database on a monthly basis for the whole sample. The different frequency required an additional aggregation step where the smoothed probabilities were averaged according to the month of each observation. The resulting series, together with the dominant regime in that month, are shown in figure 5.

Figure 5. Monthly net trade and regime probabilities



Because of the aggregation necessity, regime 1 entirely vanishes from the picture because it is never the dominant regime for a whole month. The graph illustrates that the quantity of net trade is substantially lower in months were, on average, regime 2 is dominating. Validity of the law of one price seems to hold only in those periods, when the actual quantity of trade is rather low. This result is also found when we compare the average net trade volumes in regimes 2 and 3: In the latter, the average monthly net trade quantity is 73 % higher than in the former regime.

5. Summary

In this paper, we have analysed the dynamics of wheat trade between Hungary and Germany for the period from January 2003 until September 2007. This period was characterized by rapidly changing market conditions, which in turn was reflected in varying prices, and trade volumes, between Germany and Hungary. As a result, a standard VECM was found to be incapable of providing a congruent model of the price relationship between Hungary and Germany.

A MS error correction model with three regimes was found to be a statistically superior alternative. The model seems to appropriately capture the dynamics in the price relationship. Among the regimes, one seems to capture highly unusual price drops in the Hungarian market, while on the contrary, another regime seems to relate to normal trade, corresponding to high trade volumes. This regime is the most frequent one. The most compelling regime in economic terms (because it seems to correspond very closely to the notion of the law of one price) occurs less frequently. The volume of trade is relatively low, although equilibrium deviations are reduced most quickly in this regime. It seems to be a very promising line of further research to take a more fine-grained look at the actual trade flows, and at general indicators of market conditions (e.g., market information systems, media, etc.) in order to overcome the interpretative problems because of the lack of trade quantity data at the desired frequency.

References

Bessler, D.A. – Yang, J. – Wongcharupan, M. (2003): Price Dynamics in the International Wheat Market: Modeling with Error Correction and Directed Acyclic Graphs. Journal of Regional Science, 43 (1) 1-33.

Candelon, B – Lütkepohl, H. (2000): On the reliability of Chow type tests for parameter constancy in multivariate dynamic models. Discussion paper, Humboldt University Berlin.

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, K.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. 149-173.

Ghoshray, A. (2007): An Examination of the Relationship Between U.S. and Canadian Durum Wheat Prices. Canadian Journal of Agricultural Economics, 55, 49-62

Ghoshray, A. (2002): Asymmetric Price Adjustment and the World Wheat Market. Journal of Agricultural Economics, 53 (2) 299-317

Goodwin, B. K. – Piggott, N. E. (2001): Spatial Market Integration in the Presence of Threshols Effects. American Journal of Agricultural Economics, 83(2).

Hamilton, J.D. (1989): A new approach to the economic analysis of nonstationary time series and the business cycle. Econometrica, 57, 357-384.

Gregory, A.W. - Hansen, B.E. (1996): Residual-Based Tests for Cointegration in Models with Regime Shifts, Journal of Econometrics, Vol. 70, 99-126.

Johansen, S. (1988): Statistical Analysis of Cointegrating Vectors, Journal of Economic Dynamics and Control, Vol. 12, 231-254.

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

Krolzig, H.-M. (2004): An Ox package designed for the econometric modelling of univariate and multiple time series subject to shifts in regime, Version 1.31k, <http://www.economics.ox.ac.uk/research/hendry/krolzig/msvar.html>.

Perron, P. (1997): Further evidence on breaking trend functions in macroeconomic variables, Journal of Econometrics, Vol.80, 355-385.

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, 381-397

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

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