Short- and long-run relationships between Ukrainian barley and world feed grain export prices

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ABSTRACT

Over the past decade, Ukraine has become an important player in the international feed grain market. From 2004/05 to 2012/13 it exported on average 25 percent of the total world barley annually, less than a percent lower than the largest barley exporter in the world - Australia. This research summarizes the short- and long-run barley price dynamics between Ukraine, and other major barley exporters - Australia, European Union (EU), and Canada – from 2004 to 2010. We also include U.S. corn prices to check if there is any long-run relationship between these two feed grain prices. Tests of market price cointegration (Johansen ML test and residual-based tests) and threshold error correction techniques were performed for this purpose. The results suggest that the cointegrated pairs of prices are Ukraine-Australia, Ukraine-France, Australia-Canada, and Australia-France. The estimated long-run barley price transmission elasticity is 0.71 between Ukrainian and French (a representative country of the EU) barley prices, 0.59 between Australian and Ukrainian barley prices, 0.54 between Canadian and Australian barley prices, and 0.57 between Australian and Canadian barley prices. We also found the short-term relationships between the cointegrated prices to be statistically significant. Moreover, Ukrainian barley prices were found to be weakly exogenous with regards to the Australian and French barley prices in the analyzed period, while Australian barley price is weakly exogenous with regards to the French barley price. Price adjustments in all cointegrated price series were found to be symmetric.

Key words: spatial price transmission, TAR, M-TAR, barley export prices, Ukraine, weak exogeneity.

INTRODUCTION

Over the past decade, Ukraine has become an important player in the international barley market. From 2004/05 to 2012/13 it exported on average 25 percent of the total world barley annually, which has brought Ukraine to the 2nd place in the amount of exported barley after Australia (on average 26 percent of the world barley exports per year) (USDA).

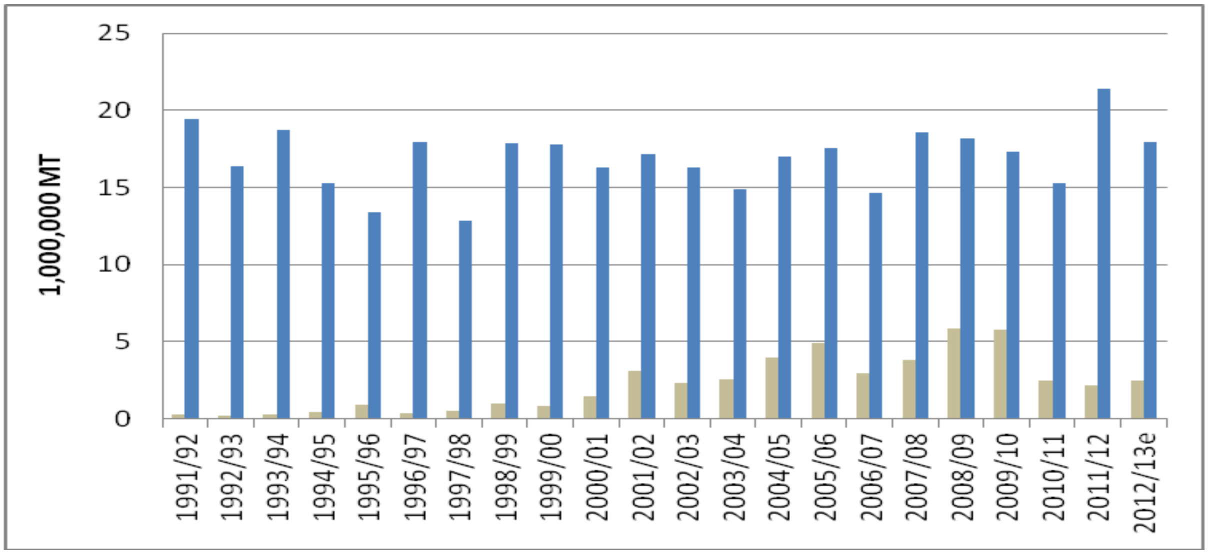

Figure 1. Ukrainian and world barley exports between 1991/92 and 2012/13

Source: USDA, January 2013

Despite the already large share of Ukrainian barley exports in the world, it could increase further due to an increase in land use and/or yields. An FAO-EBRD (2009) study suggests that up to 3 million ha in Ukraine could be added to crop production. As to yields, in the last four years Ukrainian barley yields were on average half of those in the EU-27 (USDA 2013).

Given the increasing significance of Ukrainian barley exports in the international grain market place, which also coincides with increasing commodity price volatility around the globe, it is important to learn more about the country's role in feed grain market price dynamics. The degree to which prices are transmitted to and from the region might significantly influence not only the production incentives in the domestic markets, but could also have an impact on the world market due to the large share of Ukrainian barley in the world exports. Moreover, understanding how well price is transmitted among the countries and what affects such transmission is a prerequisite for analysis of past and possible future policies and interventions. This is especially important because of Ukraine's tendency to restrict exports when world grain prices surge or domestic production falls.

The objective of this study is to analyze the short- and long-run export price dynamics between Ukraine (barley) and other major feed grain exporters – Australia (barley), European Union (barley), Canada (barley) and United States (corn). In particular, we examine the relationship between Ukrainian barley prices and the feed grain prices of the above mentioned countries using cointegration and asymmetric error correction approach. The results of the study would allow us to gain insights in how efficiently the Ukrainian barley market is integrated (if at all) with the largest feed grain exporters and analyze the possible policy issues that could stem from different levels of price transmission. Even though price relationships in spatially separated agricultural commodity markets have received a considerable amount of attention by economists in recent years, due to the growing trend towards market liberalization across the globe, no research effort has been dedicated to investigation of Ukrainian barley price dynamics. Thus, our paper is the first study that conducts a comprehensive analysis of price transmission levels (both short- and long- run) for Ukrainian barley from 2004 till 2010. Moreover, our study also contributes to the literature by estimating long-run price transmission elasticities that could be used to better link Ukrainian barley prices to other prices in modeling global barley market behavior.

The rest of the article is organized as follows: The next sections briefly highlight the conceptual framework of the study, the econometric methods to be used and the description of the data. The results of the short- and long-run price dynamics analysis and implications of the study are contained in the final sections.

ECONOMETRIC METHODS

The notion of price cointegration and, as a result, of market integration lies in the Law of One Price (LOP), according to which in an efficient market the uniform goods must have only one price once transportation costs are accounted and assuming the absence trade restrictions (Isard 1977). This is the strong version of the LOP. It holds on the condition of spatial arbitrage, which suggests that if the prices of two identical goods have different prices in different locations, the higher prices will attract the arbitrageurs to take advantage of the existing profits until the point when the prices equalize across the different locations. Thus, in the short-run prices can deviate from one another, but in the long run they will be the same after accounting for transportation costs.

In reality, however, there are a number of factors that could affect the efficiency of markets and/or price relationships between different goods, such as, transaction costs, market power, exchange rates, quality differences, etc. This results in a failure of most empirical tests to support the hypothesis of the LOP, which "might depend both on the strong assumptions underpinning it and on the inherent features of the empirical models used" (Listorti 2008). Therefore, most economists tend to focus on testing market integration (or a weaker version of LOP), rather than adherence to the strong version. In particular, under the weaker condition of LOP two spatially separated markets are considered to be integrated for a particular good if there is a long-run relationship between the prices for this good in different markets.

There are a large number of empirical models used for spatial price analysis, however time series analysis and, in particular, cointegration models are the most widely used for the analysis of price transmission. Their popularity can be explained by a number of benefits their use provides. First of all, cointegration models allow analyzing both short- and long-run price dynamics. Second, they can provide reliable results when the only data available are prices. However, one needs to remember that the interpretation of such results needs to be conditional on the assumption that there exist continuous and unidirectional trade linkages among the analyzed countries. Therefore, the conclusions need to be carefully drawn with the specifics of the particular market in mind. For example, the absence of cointegration might not necessarily be a guarantee of lack of market integration, but of a need to research other factors that could affect price cointegration and market efficiency. Finally, one more benefit of using cointegration models is that they do not require the assumption of exogeneity for the analyzed price series.

In our analysis in order to test for the long-run relationship between Ukrainian barley and US corn, Canadian, EU, and Australian barley prices we used both the Engel- Granger cointegration procedure (the primary one) and the Johansen maximum likelihood (ML) cointegration test (for assessing the robustness of the results). We ran the Johansen ML test both on the multiple series to estimate the total number of cointegrating relationships and on different pairs of the barley or corn price series. Despite some shortcomings mentioned before, both of the methods are commonly used to test for cointegration in commodity markets (see Ghosray and Lloyd 2003; Listorti 2008).

To be able to test two price series for the cointegrating vector, we first need to confirm the presence of a unit root within each series, which indicates the series is non-stationary. For this purpose, the Augmented Dickey-Fuller (ADF), Philips-Perron (PP), and Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) tests were conducted.

The main objective of the ADF (Dickey and Fuller 1979) unit-root test is to test the null hypothesis that $\beta_1 = 0$ (i.e. Ho: series contains a unit root) in the following equation:

$$\Delta P_t^{UKR} = a_0 + a_1 t + \beta_1 Y_{t-1}^{UKR} + \sum_{j=1}^{p} \gamma_j \Delta Y_{t-j}^{UKR} + \varepsilon_t \,, \quad (1)$$

where P^{UKR} is the Ukrainian barley price in logarithms, and Δ denotes first difference. Similar equations were run for the Australian, Canadian, EU barley and US corn log-price series. The above equation includes a time trend, represented by *t*. In those cases when we checked for the unit roots in the time series with an intercept only (i.e. excluding a time trend), *t* was not included.

The number of lags was estimated by minimizing the Schwarz Bayesian Criteria (SBC) starting with the 12 lags in the initial regressions because it is monthly data. The correct choice of the lag length is important. If the number of lags is too small, the error terms will be serially correlated and the results of the tests may be biased. On the contrary, the more lags are added, the more degrees of freedom are lost.

The Philips -Perron test statistic is similar in interpretation to the ADF ones and usually provides the same results. The advantage of the PP test statistic, however, is that it “incorporates an automatic correction to the ADF procedure to allow for autocorrelated residuals” (Brooks 2002).

The KPSS unit root tests (Kwiatkowski et al. 1992) were run as a robustness check of the results obtained from the ADF and PP tests. As was mentioned before, the null hypotheses of both ADF and PP tests assume non-stationarity of the series, which results in a low power of these tests to reject the null, unless there is strong evidence of the stationarity. This might result in Type II errors. On the contrary, the KPSS test's null hypothesis is that the data is stationary. Due to these differences in the designs of the tests, KPSS is a good complement to the ADF and PP unit root tests. For example, if the ADF and PP tests fail to reject the null hypothesis, while KPSS rejects its null, strong evidence of the unit root presence can be assured. If however, one of the tests does not support the evidence of another, further investigation of the series is needed (Cheung et al 1994).

Two non-stationary series that are integrated of the same order are cointegrated if they have a long-run relationship and a linear combination of the series is stationary, even if they diverge in the short run. Moreover, if series are cointegrated it implies that the weaker condition of the LOP holds.

As was mentioned earlier, Johansen's cointegration test (Johansen 1988) is commonly used to test for the presence of cointegrating vectors. To obtain the test results, we first specify the VAR(k) model, where k is the number of lags:

$$P_t = \alpha_0 + a_1 t + A_1 P_{t-1} + \cdots + A_p P_{t-p} + \varepsilon_t, \qquad (2)$$

where P_t is an *n x 1* vector of prices, and A_p is the matrix of the coefficients to be estimated. This equation is further converted into the following vector error correction model:

$$\Delta P_t = \alpha_0 + a_1 t + \prod P_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta P_{t-i} + \varepsilon_t, \qquad (3)$$

where Δ denotes first difference, $a_1 t$ is a deterministic trend, Γ_i represents the dynamic effects, while Π captures the long-run effects of the analyzed series. The goal of the Johansen ML test is to estimate the rank of the Π matrix, which represents the number of cointegrating relationships.

The major difference between the Johansen ML and Engle-Granger methods is that they require different model assumptions. The first one requires a normality assumption, while the latter one is insensitive to the distribution assumption. Therefore, one of the benefits of using the Engle-Granger method is in its relative efficiency over the Johansen ML test if normality does not hold. As to the benefits of using the Johansen ML method, it allows obtaining more than one cointegrating relationship, and this is the major reason why it was used in our study.

The residual-based test for cointegration, Engle- Granger (1987) procedure, consists of two steps. First, the long run relationship between the pairs of export log-prices is estimated as seen in the example of the relationship between Ukrainian and Australian barley prices:

$$P_t^{UKR} = \beta_0 + \beta_1 P_t^{AUS} + \varepsilon_t \qquad (4),$$

where P_t^{UKR}, P_t^{AUS} are prices of Ukrainian and Australian barley, respectively. β_0 accounts for the transfer costs, β_1 stands for the price transmission elasticity, and ε_t is the error term.

Second, we test whether $\gamma_1 = 0$ (i.e. unit root is present and the residuals are non-stationary) in the following regression:

$$\Delta\bar{\varepsilon_t} = \gamma_1 \bar{\varepsilon}_{t-1} + \sum_{i=1}^{p} \gamma_{i+1} \Delta \bar{\varepsilon}_{t-i} + \omega_t \quad (5)$$

where ω_t is the white noise term, and $\bar{\varepsilon_t}$ is the residual obtained from the long-run equilibrium equation (4). As previously stated, the number of lags is selected by minimizing the SBC and making sure that errors are not serially correlated. Rejecting the null would mean that analyzed barley/corn prices are cointegrated, i.e. they move together in the long-run. To test for autocorrelation of the residuals the Breusch-Godfrey (Breusch 1979) test was used.

If two series are cointegrated, then the OLS estimators are superconsistent, which implies that they converge at a faster than normal rate, and they can be used to characterize the series' behavior. We used both the residual-based and Johansen ML tests for the robustness check.

The studies by Enders and Granger (1998) and Enders and Siklos (2001), however, suggest that the tests for cointegration could provide inconsistent results if the price adjustment is asymmetric. Therefore, they suggested a modification to equation (5) to test for asymmetric price transmission, which is known as the TAR (threshold autoregressive) model:

$$\Delta\bar{\varepsilon}_t = I_t\gamma_1\bar{\varepsilon}_{t-1} + (1 - I_t)\gamma_2\bar{\varepsilon}_{t-1} + \varphi_t, \qquad (6),$$

Where $\Delta\bar{\varepsilon_t}$ is the first difference of the error term from (4), $\bar{\varepsilon}_{t-1}$ is lagged error term from (4) lagged for one time period, γ_1 and γ_2 are the adjustment rates, $I_t = \begin{cases} 1 \ if & \bar{\varepsilon}_{t-1} \geq \tau \\ 0 \ if & \bar{\varepsilon}_{t-1} < \tau \end{cases}$ (7), and τ is equal to zero. Equation (6) can also be specified with the additional lags of $\Delta\bar{\varepsilon_t}$ to control for serial correlation. In our study the number of lags was based on minimizing the SBC. Threshold equal to zero implies that adjustment is equal to $\gamma_1\bar{\varepsilon}_{t-1}$ if $\bar{\varepsilon}_{t-1}$ is above long-run equilibrium value and $\gamma_2\bar{\varepsilon}_{t-1}$ if $\bar{\varepsilon}_{t-1}$ is below long-equilibrium (Enders et al. 2001).

The TAR model has the purpose of analyzing any "deep" movements in the series (Enders and Granger, 1998). In order to capture any "steep" variations in ε_t, Enders and Siklos (2001) suggested an alternative, M-TAR (momentum threshold autoregressive) model. Unlike TAR, where I_t depends on the levels of the error term (ε_t), in the M-TAR the value of the indicator function I_t depends on the change in ε_{t-1} in the previous period. Accordingly, equation (7) is modified in the following way:

$$I_t = \begin{cases} 1 \; if \; \Delta\bar{\varepsilon}_{t-1} \geq \tau \\ 0 \; if \; \Delta\bar{\varepsilon}_{t-1} < \tau \end{cases} \quad (8), \text{ where}$$

as previously, τ is equal to zero. M-TAR model is useful when it is expected that the series ε_t exhibits more momentum with regards to either increase or decrease in price.

For both TAR and M-TAR models, the first step is to check (and in the case of our paper, to confirm) that the analyzed series are cointegrated. To do so, the null hypothesis H_0: $\gamma_1 = \gamma_2 = 0$ of no cointegration is tested. Since the F statistic for the above null hypothesis has a non-standard distribution, the Φ-statistic is used instead (see Enders and Granger 2001). If the null that $\gamma_1 = \gamma_2 = 0$ is rejected, we can conclude that the series are cointegrated and proceed with the test for the symmetric price adjustment. To do so, the null hypothesis of symmetric adjustment H_0: $\gamma_1 = \gamma_2$ is tested. Standard F-statistics can be used to test this hypothesis. If we fail to reject the null, we can conclude that price adjustment is symmetric. Rejecting the null, however, would suggest that the series responds differently to whether the departure from the long-run equilibrium is increasing or decreasing. If, for example, $|\gamma_1| < |\gamma_2|$, this would suggest that increases in the price tend to persist, while the decreases are transmitted more rapidly back to the long run equilibrium.

Testing for asymmetric adjustment is important for several reasons. First, if the long run relationship between two series is found to be asymmetric, the results of the cointegration tests described earlier may provide misleading results (Frey and Manera 2007), and such tests need to be adjusted to account for the asymmetry. Second, if the price transmission is asymmetric it might have important implications for the consumer and producer welfare effects and should be taken into account by policy makers (Awokuse and Wang 2009). Finally, testing for the asymmetric price transmission provides researchers as well as policy makers with insights into the potential market inefficiencies that could be further analyzed.

The literature suggests that both TAR and M-TAR models are commonly used methods to test for asymmetric price transmission. For the examples of the TAR and M-TAR methods applied to the analysis of the agricultural markets see Sephton (2011), Awokuse and Wang (2009), Ghoshray (2007), and Abdulai (2000).

If two price series are cointegrated and the price adjustment is symmetric, their short-run dynamics can be analyzed by using an error correction model of the following form (using the example of Ukraine and Australia):

$$\Delta P_t^{UKR} = a_0 + a_1\bar{\varepsilon}_{t-1} + \delta_0 \Delta P_t^{AUS} + \sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{UKR} + \sum_{j=1}^{p} \theta_j \Delta P_{t-j}^{AUS} + \mu_t \quad (9.1) \text{ and}$$

$$\Delta P_t^{AUS} = a_0 - a_1\bar{\varepsilon}_{t-1} + \delta_0 \Delta P_t^{UKR} + \sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{AUS} + \sum_{j=1}^{p} \theta_j \Delta P_{t-j}^{UKR} + \mu_t$$

where ΔP_t^{UKR} and ΔP_t^{AUS} are vectors of the first differences of log prices for Ukraine and Australia, $\bar{\varepsilon}_{t-1}$ is the lagged residual from (4), μ_t is the error term, and the scalar a_1 represents the short-run adjustment speed of the dependent variable to the long-run steady state (Baffes 2003). Its sign is expected to be negative. Equation 9 can be specified with additional lags ($\sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{AUS}$ and $\sum_{j=1}^{p} \theta_j \Delta P_{t-j}^{UKR}$) to deal with autocorrelation which might be present in the error term. The appropriate lag length was selected by minimizing the SBC, and using the Breusch-Godfrey test. The deterministic trend was not included in (9.1 and 9.2) since it was found to be statistically insignificant.

If the short-run adjustment coefficient is to be found statistically insignificant, this would suggest that series are weakly exogenous. If price is found to be weakly exogenous, the implication is that the price evolves independently and other prices adjust to maintain long-run equilibrium.

If the null hypotheses in the TAR or/and M-TAR models are rejected, the ECM needs to be modified to account for asymmetry in the price transmission. One of the examples of such an ECM is the threshold error-correction model:

$$\Delta P_t^{UKR} = a_0 + \rho_1 I_t \bar{\varepsilon}_{t-1} + \rho_2 (1 - I_t) \bar{\varepsilon}_{t-1} + \sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{UKR} + \sum_{j=1}^{n} \theta_j \Delta P_{t-j}^{AUS} + \mu_t \quad \text{(10.1) and}$$

$$\Delta P_t^{AUS} = a_0 - \rho_1 I_t \bar{\varepsilon}_{t-1} + \rho_2 (1 - I_t) \bar{\varepsilon}_{t-1} + \sum_{i=1}^{p} \delta_i \Delta P_{t-i}^{AUS} + \sum_{j=1}^{n} \theta_j \Delta P_{t-j}^{UKR} + \mu_t \quad \text{(10.2),}$$

where ρ_1 and ρ_2 represent the speed of adjustment depending on whether $\bar{\varepsilon}_{t-1}$ or $\Delta\varepsilon_{t-1}$ is above or below the threshold. In this study we focus on testing long run asymmetry only, however, one should note that the model in (10.1 and 10.2) could be further modified to incorporate short run asymmetries.

DATA

In this empirical investigation monthly FOB prices for Ukrainian Feed Barley (Black Sea), Australian Feed Barley (Southern States), French Feed Barley (Rouen), Canadian Malting Barley (Thunderbay) and U.S. No 3. Yellow Corn (Gulf) are used. The time span of our analysis is from November 2004 till October 2010. We assumed the French barley price to be representative of the EU barley price, since it is the largest exporter of feed barley in the EU (Eurostat). Canadian malting barley prices serve as a proxy for the Canadian feed grains, though we acknowledge the possibility that the difference in the quality of Ukrainian and Canadian barley might negatively affect the cointegration results between these two countries. The series were obtained from the International Grains Council and HGCA.

A visual inspection of the graph of the analyzed series in USD per ton suggests that in general all the series tend to move together over the analyzed period (figure 2), however, the magnitude of these similarities differs. Ukrainian, French and Australian barley prices seem to move in the most similar manner over the investigated time span. U.S. corn prices are consistently lower than the barley ones, though the overall direction of the movements coincide with the barley series.

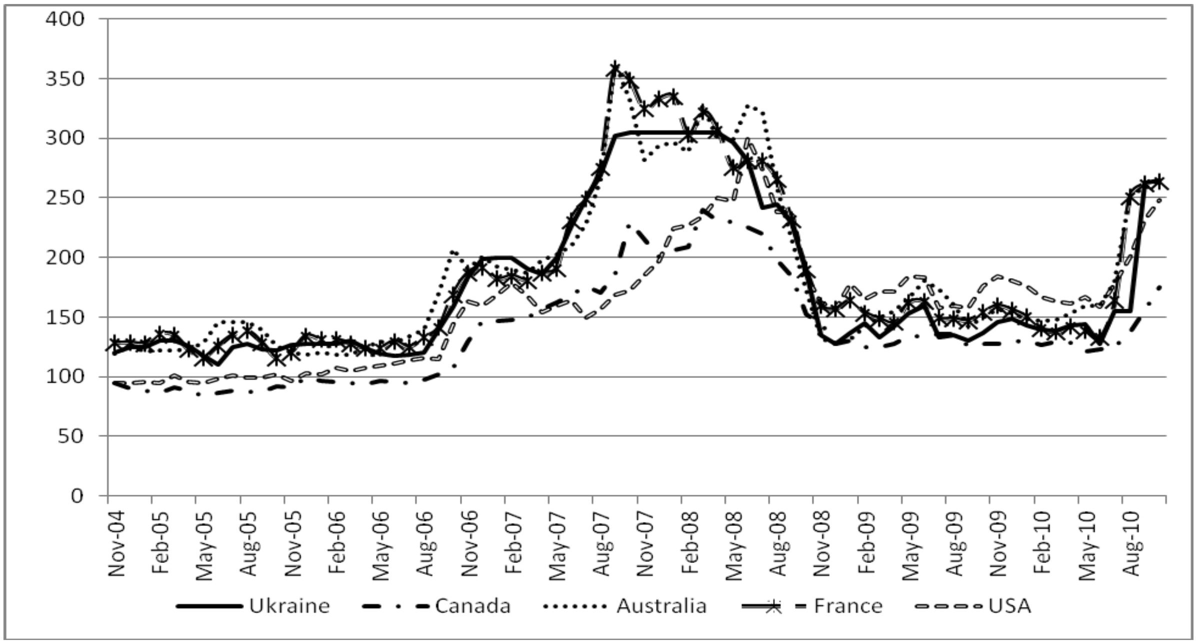

Figure 2. Comparison of export prices at different export points, $ per ton

Source: International Grain Council, HGCA 2012

EMPIRICAL RESULTS

Prior to the model estimation we determined the order of integration of the analyzed series by using the unit-root tests. All three tests (ADF, PP, and KPSS) supported the evidence of the unit-root presence in the series. Thus, the tests were re-run on the series after they were differenced in log levels. The results provided in the table 3 show that all the differenced series are stationary. This leads to the conclusion that the price series of Ukraine, Canada, Australia, EU and US are I(1). The tests were run for the cases when trend is present, and when it is absent. In case of the KPSS test results for the first-differences series of Ukraine and Canada, the presence of trend makes test results significant at 10% level; however, these results are on border to be insignificant. And since all other unit root tests do not differentiate between the presence or absence of a trend, we conclude that trend inclusion does not affect the outcome.

Table 2. Results of the unit root tests in levels[a]

		ADF		PP		KPSS	
	# of lags	w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
Ukraine	1	-1.84	-1.96	-1.25	-1.36	0.43*	0.29**
Australia	1	-1.95	-2.09	-1.51	-1.61	0.48**	0.28**
France	4	-1.66	-1.85	-1.46	-1.60	0.48**	0.30**
Canada	1	-1.39	-1.5	-1.08	-1.19	0.79**	0.35**
USA	1	-1.25	-1.9	-1.11	-1.72	1.28**	0.31**

[a]Asterisks denote levels of significance (* for 10 percent , ** for 5 percent). The 5% and 10% critical values for ADF and PP tests with a drift are -2.90 and -2.59 respectively; for the tests with a drift and a trend are -3.47 and -3.16 respectively. Critical values were obtained from MacKinnon (1991). The 5% and 10% critical values for the KPSS test in levels are 0.463 and 0.347 respectively; for the KPSS tests with a trend they are 0.146 and 0.119 respectively.

Table 3. Results of the unit root tests using first difference[b]

		ADF		PP		KPSS	
	# of lags	w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
Ukraine	4	-3.04**	-2.96**	-4.83**	-4.80**	0.12	0.12*
Australia	1	-4.89**	-4.85**	-5.45**	-5.41**	0.09	0.09
France	4	-5.42**	-5.38**	-5.51**	-5.47**	0.11	0.117
Canada	1	-3.97**	-3.93**	-5.62**	-5.58**	0.14	0.12*
USA	1	-4.78**	-4.73**	-6.92**	-6.87**	0.08	0.08

[b]Asterisks denote levels of significance (* for 10 percent , ** for 5 percent). The 5% and 10% critical values for ADF and PP tests with a drift are -2.90 and -2.59 respectively; for the tests with a drift and a trend are -3.47 and -3.16 respectively. Critical values were obtained from MacKinnon (1991). The 5% and 10% critical values for the KPSS test in levels are 0.463 and 0.347 respectively; for the KPSS tests with a trend they are 0.146 and 0.119 respectively.

Concluding that the analyzed series are I(1) allowed us to proceed to the cointegration tests. We start with the Johansen ML test on all the series of interest to test for the total number of the long-run co-integrating vectors. In order to do so, the appropriate order of VAR is first established for each series, using the SBC. It was equal to 2.

The results that are provided in table 4 suggest that we reject the null hypothesis of no cointegration, i.e. r = 0, where r is the number of cointegrating relationships. We also reject H_0 that there is only one cointegrating relationship (r =1). Since, however, we fail to reject the null hypothesis that r = 2, we conclude that there are two or more distinct long-run relationships among the five series.

Table 4. Cointegration rank test using trace[c]

Ho(Rank=r)	H1(Rank>r)	Trace	5% CV
0	0	87.99**	75.74
1	1	55.88**	53.42
2	2	26.31	34.80
3	3	12.22	19.99
4	4	3.82	9.13

[c]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent).

In order to find out which pairs of series are integrated, both Johansen's ML and Engle- Granger cointegration tests were run on the pairs of series. Engle- Granger tests suggest that the cointegrated pairs of series are Ukraine-Australia, Ukraine-France, Australia-Canada, and Australia-France (table 5).

Table 5. Engle- Granger cointegration tests for barley/corn price series of interest[d]

Pair of series	# of lags	ADF	PP	KPSS
Ukraine-France	1	-4.44**	-4.88**	0.21
Ukraine-Canada	3	-1.72	-2.51	0.72**
Ukraine-Australia	1	-4.67**	-4.23**	0.16
Ukraine-USA	1	-1.66	-1.56	0.59**
France-Australia	2	-3.56**	-3.54**	0.09
France- Canada	2	-2.54	-3.00	0.51**
France-USA	1	-1.95	-1.77	0.51**
Australia-Canada	1	-3.47**	-3.15*	0.32
Australia-USA	2	-2.22	-2.07	0.50**
Canada-USA	1	-1.94	-2.15	0.52**

[d]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent). The 5% and 10% critical values for tests with a drift are -3.37 and -3.07 respectively. Critical values were obtained from MacKinnon (1991). The 5% and 10% critical values for the KPSS test in levels are 0.463 and 0.347 respectively; for the KPSS tests with a trend they are 0.146 and 0.119 respectively.

The Johansen ML pair wise tests confirm cointegration of Ukraine-Australia, Ukraine-France, and Australia-France pairs of barley prices, however, not the Australia-Canada relationship (Table 6). The difference in the outcomes from different cointegration tests might be attributed to the different sets of assumptions that are used when constructing both tests. Since Engle-Granger tests are less sensitive to the lag selection, we set it as our priority test; therefore, we conclude that Australian and Canadian barley series are also cointegrated. This would be further confirmed by the TAR/M-TAR tests for this pair of prices.

Table 6. Johansen ML Pairwise cointegration tests for barley/corn price series of interest[e]

Pairs of series	Ho(H1)	Trace	5%CV
Ukraine-France	r=0((r>0)	21.43**	19.99
	r=1(r>1)	4.10	9.13
Ukraine-Canada	r=0((r>0)	17.78	19.99
	r=1(r>1)	3.18	9.13
Ukraine-Australia	r=0((r>0)	21.16**	19.99
	r=1(r>1)	4.14	9.13
Ukraine-USA	r=0((r>0)	9.79	19.99
	r=1(r>1)	3.25	9.13
France-Australia	r=0((r>0)	22.78**	19.99
	r=1(r>1)	3.63	9.13
France-Canada	r=0((r>0)	14.71	19.99
	r=1(r>1)	2.79	9.13
France-USA	r=0((r>0)	9.87	19.99
	r=1(r>1)	3.18	9.13
Australia-Canada	r=0((r>0)	18.30	19.99
	r=1(r>1)	2.58	9.13
Australia-USA	r=0((r>0)	12.02	19.99
	r=1(r>1)	3.23	9.13
Canada-USA	r=0((r>0)	7.92	19.99
	r=1(r>1)	2.76	9.13

[e]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent).

Since Ukraine-Australia, Ukraine-France, Australia-Canada, and Australia-France series are co-integrated, the results of the regressions that analyze the relationships between them are consistent (see equation 4). Thus, β_1 can be considered as the long-run price transmission elasticity. The results estimate that the long-run price elasticity is 0.71 between Ukrainian and French barley series, 0.59 between Australian and Ukrainian barley series, 0.54 between Canadian and Australian barley series, and 0.57 between Australian and Canadian barley series. The price transmission elasticity indicates the percentage change in the price of one country in response to a one-percent change in the world market. It is directly related to trade liberalization (Listorti 2008; Thompson 1999), since higher levels of trade liberalization contribute to greater price transmission elasticities.

The next step is to proceed with testing for asymmetric price transmission in the pairs of prices that were found to be cointegrated. In the cases of all four pairs of prices, we reject the null of no cointegration (H_0: $\gamma_1 = \gamma_2 = 0$), and confirm our previous results of long-run relationships between Ukraine-France, Ukraine-Australia, Australia-Canada, and Australia-France[1] (table 7). The F-statistic estimates for both TAR and M-TAR models for all pairs of series considered suggest that we cannot reject the null of the symmetric price transmission (tables 7 and 8). Therefore, we conclude that the price transmission between the four pairs of prices is symmetric.

Table 7. TAR model parameter estimates[f]

	Ukraine - France	Ukraine-Australia	Australia-Canada
Variable	**Parameter estimate**	**Parameter estimate**	**Parameter estimate**
γ_1	-0.93 (-4.25)**	-0.54 (-2.16)**	-1.06 (-6.78)**
γ_2	-1.31 (-6.69)**	-0.97 (-4.70)**	-0.90 (-4.68)**
$H_0: \gamma_1 = \gamma_2 = 0 (\Phi)$	25.43**	11.37**	33.91**
$H_0: \gamma_1 = \gamma_2 \ (F)$	2.35[0.13]	2.55 [0.12]	0.42 [0.52]

[f]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent). The 10% and 5% significance level critical values are 3.79 and 4.64 respectively. t-values are stated in parenthesis. The values in the square brackets denote the p-values.

Table 8. M-TAR model parameter estimates[g]

	Ukraine - France	Ukraine-Australia	Australia-Canada
Variable	**Parameter estimate**	**Parameter estimate**	**Parameter estimate**
γ_1	-1.15 (-6.52)**	-0.89 (-3.26)**	-0.97(-4.67)**
γ_2	-0.78 (-3.88)**	-0.39 (-1.84)**	-1.02 (-5.30)**
$H_0: \gamma_1 = \gamma_2 = 0 (\Phi)$	28.54**	6.39**	23.33**
$H_0: \gamma_1 = \gamma_2 \ (F)$	2.00 [0.16]	2.41 [0.13]	0.05[0.83]

[g]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent). The 10% and 5% significance level critical values are 4.11 and 5.02 respectively. t-values are stated in parenthesis. The values in the square brackets denote the p-values.

[1] Preliminary results for the TAR and M-TAR models for the Australia-France pair are subject to further investigation due to the presence of serial correlation in the residuals. For this reason they are not included here.

The absence of asymmetric price transmission allows us to proceed with the construction of a simple error-correction model for the cointegrated series to analyze the short term dynamics between them. The result show that Ukrainian barley price are weakly exogenous with regards to French and Australian prices, which implies that Ukraine price evolves independently, while French and Australian prices adjust after the change in the Ukrainian price to maintain long-run equilibrium. About 39 percent of the adjustment takes place within one month in the case of the Ukraine-France relationship, and approximately 32 percent for the Ukraine-Australia pair of prices. Australian prices are also found to be weakly exogenous with regards to the French prices, which positions France as a clear price follower in the international barley market. After a change takes place in the Australian price, during the first month French price adjusts back to the equilibrium by 26 percent.

For the pair Canada-France, none of the series exhibit weak exogeneity. However, the coefficients of the error-correction terms in the corresponding equations are found to be statistically significant (table 9). This suggests that these series adjust to the long-run equilibrium after the change in the corresponding price with 15-21 percent of price adjustment happening in one month.

Table 9. Error-correction model parameter estimates[h]

	# of lags	Speed of adjustment, α_1	Test F-value
Ukraine-France	2 ; 2	-0.13 (-0.70)	9.60**
France-Ukraine	1 ; 1	-0.39 (-2.14)**	6.04**
Australia-Ukraine	1 ; 0	-0.32 (-2.61)**	10.49**
Ukraine-Australia	1 ; 1	-0.04(-0.56)	9.29**
Australia-France	1 ; 1	-0.18(-1.33)	4.92**
France-Australia	1 ; 1	-0.26 (-2.13)**	6.69**
Australia-Canada	1 ; 0	-0.21 (-2.20)**	9.30**
Canada-Australia	1 ; 1	-0.15 (-2.74)**	18.42**

[h]Asterisks denote levels of significance (* for 10 percent, ** for 5 percent).

CONCLUSIONS AND IMPLICATIONS

We investigated the long- and short-run dynamics between Ukrainian and Australian, Canadian and EU barley prices as well as US corn prices using monthly FOB data November 2004 till October 2010. The results suggest that the cointegrated pairs of prices are Ukraine-Australia, Ukraine-France, Australia-Canada, and Australia-France. The estimated long-run barley price transmission elasticity is 0.71 between Ukrainian and French (a representative country of the EU) barley prices, 0.59 between Australian and Ukrainian barley prices, 0.54 between Canadian and Australian barley prices, and 0.57 between Australian and Canadian barley prices.

We also found the short-term relationships between the cointegrated prices to be statistically significant. The error correction model results showed that about 39 percent of the adjustment back to the equilibrium takes place within one month in the case of the Ukraine-France relationship, and approximately 32 percent for the Ukraine-Australia pair of prices. For Australia-Canada and Australia-France pairs of prices, the first month adjustment is about 20 percent depending on the pair under consideration. Such slow adjustments for all the barley series suggests that there are inefficiencies present in the markets of the analyzed countries that need to be studied further. The results of the TAR and M-TAR model suggested that shocks between the cointegrated series are transmitted symmetrically. This lets us conclude that the weaker version of the LOP does hold in case of the Black Sea barley market, and that and Ukrainian barley price series are closely following the world barley market dynamics. Moreover, Ukrainian barley prices were found to be weakly exogenous with regards to the Australian and French barley prices in the analyzed period. This suggests that Ukrainian barley price evolves independently and, thus, exhibits price leadership behavior in the international barley market.

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