

THREE ESSAYS ON BLACK SEA GRAIN MARKETS

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DEDICATION

I would like to dedicate this work...

To my grandparents for always believing in me;

To my parents for giving me both strong family roots and large wings to fly high above;

To my husband for supporting each step I make, for his ability to cheer me up no matter what, and above all, for making me happy every minute of my life;

To the Schroeders and Scheperles for becoming my family and making me feel at home;

To Diane for always being my ray of sunshine and calling me Dr. Goychuk well before I officially earned that title;

To Winston for unexpectedly turning my world upside down;

To my friends, my relatives, and my country for shaping me into what I am now.

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THREE ESSAYS ON BLACK SEA GRAIN MARKETS

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ABSTRACT

Russia and Ukraine (often referred to as the Black Sea region) have recently emerged as major world grain exporters. Just in the 2011/12 marketing year, according to the USDA (2013) they were among the top ten global suppliers of wheat (27 mln. tons), of corn (17.2 mln. tons), and of barley (5.8 mln. tons). The conquest of the world markets, however, did not come with a similar openness in the approach of the Ukrainian and Russian governments to policy implementation. Until recently it has still been rather dominated by frequent and rather ad hoc policy interventions in the grain markets and especially export markets.

The purpose of this dissertation is to analyze the efficiency of Black Sea grain markets within the context of highly regulated markets. The first part of Essay 1 summarizes the short- and long-run wheat price dynamics between these two countries, and other major wheat exporters - United States, European Union (EU), and Canada – from 2004 to 2010. Tests of market price cointegration (Johansen ML test and residual-based tests) as well as threshold error correction techniques were performed for this purpose. The results suggest that Russian wheat prices were cointegrated with EU (France was considered a representative country of the EU) and US wheat prices but not with Canadian wheat prices. Ukrainian wheat prices were found to be cointegrated with French wheat prices only. The estimated long-run wheat price transmission elasticities were estimated to be equal to 1.04 between Russian and French wheat prices, 1.16

between Russian and US wheat prices, and 1.05 between Ukrainian and French wheat prices. We also found the short-term relationships between the cointegrated series to be statistically significant.

In the second part of Essay 1, the focus is on the short- and long-run barley price dynamics between Ukraine, and other major barley exporters (Australia, EU, and Canada) from 2004 to 2010. U.S. corn prices were also included with the purpose of checking if there is any long-run relationship between these two feed grain prices. Tests of market price cointegration (Johansen ML and residual-based tests) as well as threshold error correction techniques were performed for this purpose. The cointegrated pairs of prices were Ukraine-Australia, Ukraine-France, Australia-Canada, and Australia-France. The estimated long-run barley price transmission elasticity was 0.71 between Ukrainian and French barley prices and 0.59 between Australian and Ukrainian barley prices. Moreover, Ukrainian barley prices were found to be weakly exogenous with regards to the Australian and French barley prices in the analyzed period.

The second essay analyzes price transmission along the Ukrainian wheat supply chain from January 2005 until December 2012 and identifies potential market and policy failures that could lead to different levels of price transmission. In our analysis we relax the assumptions of linearity and symmetric adjustment by extending the traditional cointegration models, such as Engle and Granger (1987) procedure and the Johansen Maximum Likelihood method (1988), with the Bai and Perron (2003) structural break test and threshold autoregressive (TAR and M-TAR) models, respectively. The overall results suggest that long run relations hold between world wheat price and Ukrainian wheat farm price, as well as between farm price and flour price. Results of the structural

break tests reveal that during the times of excessive government interventions in the grain export markets, the long-run price transmission between world and farm prices significantly decreases. Price transmission was also found to be asymmetric between farm and flour prices, suggesting that Ukrainian millers tend to exhibit market power and not pass on decreases in wheat prices to the bakers.

Essay 3 investigates the development of price volatility in the Ukrainian wheat market from 2005 till 2012 within a dynamic conditional correlation GARCH framework. The results indicate that the export controls in Ukraine have not significantly reduced price volatility on the domestic wheat market as was suggested by the Ukrainian policymakers. On the contrary, our findings suggest that the multiple and unpredictable interference of the Ukrainian government on the wheat export market coincided with increased market uncertainty and pronounced additional price volatility in the market. The estimates of the conditional correlations show that interdependency between Ukrainian and world wheat markets volatility is very low while volatility transmission between two markets is statistically insignificant. This further suggests that increased volatility in the domestic wheat market was caused by internal rather than external factors.

I. INTRODUCTION

According to FAO projections, by 2050 the world population will increase to 9.1 billion people (FAO 2009). Such an increase in population along with the rising incomes around the globe would cause the world demand for food to increase by 70% between 2009 and 2050. In particular, demand for cereals, for both food and feed uses, are projected to reach about 3 billion tons by that time. This would require the annual production of cereals to increase by about 25%.

Russia and Ukraine (often referred to as the Black Sea region), along with South American and African countries, are often mentioned as a region from which large increases in grain production would come. Both countries have emerged as major world grain exporters within the last ten years. Just in the 2011/12 marketing year, according to the USDA (2013) Ukraine and Russia were among the top ten global suppliers of wheat (27 mln. tons), of corn (17.2 mln. tons), and of barley (5.8 mln. tons). Such a surge in grain production and exports can be attributed to a number of factors. After the collapse of the Soviet Union both countries experienced a significant drop in agricultural production (USDA 2013). This was caused by the dismantling of the entire centrally-planned economic system and later by the privatization and restructuring processes that took place in the sector. As a result, it took some time for both Russia and Ukraine to recover their agricultural production. More specifically, grain production declined about 40 percent from 1987 to 2000; however, it almost doubled during the next decade, and starting from 2009/10 it has continuously exceeded pre-transition levels. Another factor that has contributed to the growth of grain exports from the region has been a gradual

decrease in the land area used for fodder crops, attributable to the decline in livestock numbers. For the same reason there has been a decline in the use of major crops for domestic feed. Thus, as grain production has recovered from the post-Soviet collapse and domestic feed use has declined, there has been an increase in the exports of all major Ukrainian and Russian grains starting from the mid-1990s. Wheat and corn have seen the largest increases. Starting from 2008, the Black Sea region was exporting on average 23 million tons of wheat per year, compared to 2.2 million tons per year during the 1990s. Corn exports increased from about 250 thousand tons in the 1990s to 10 million tons per year in 2008-2012 period. Barley exports increased from 420 thousand tons in 1990 to a 9.5 million tons record in 2008.

Despite the already large share of Black Sea grain exports in the world, it could further rise due to increases in the land use and/or yields. The FAO-EBRD (2009) study states that up to 6 million hectares (ha) of abandoned land in Russia and up to 3 million ha in Ukraine could be added to grain production. Yield statistics also suggest that there is strong potential for increased production of barley, corn and wheat if Black Sea yields even partially close the gap with the averages of the EU and the U.S. For example, average barley yields in Ukraine are about two-thirds the level of yields in the U.S. and are about half the level of average EU yields. The situation with wheat and corn yields indicates even more potential for improvement. Ukrainian corn yields are about 60 percent of those in the EU-27, and less than half the yields of U.S. farmers. A similar situation exists with the wheat yields. While Ukrainian and U.S. wheat yields do not differ much on average, EU-27 wheat yields are twice as high as those that are seen in Ukraine.

It cannot be expected that Ukrainian or Russian farmers will have the levels of government policy support that exist in the EU and the U.S., but given the relatively strong world prices that are projected to remain well above the pre-2005 levels (FAPRI 2013), the economic incentive for improved inputs and management practices will continue to be strong. The question, however, is whether the farmers from Black Sea region can take advantage of the growing world grain prices. Unfortunately, both Russian and Ukraine have been plagued by erratic and often unfavorable government policies. During the financial crisis in 2006-08 and later in 2010-11, both countries introduced a variety of grain export restrictions that took the form of bans, quotas, and tariffs. Being disruptive to the markets by themselves, these trade policy interventions were also accompanied by a dramatic increase in market uncertainty as our research shows. For example, the export quotas were implemented on short notice or in some cases were not implemented at all after being announced. The size of the quotas was readjusted multiple times a year. And as anecdotal evidence suggests, quota distribution came with quite an increase in rent-seeking behavior among exporters. Such an unstable environment distorts farmers' incentives locally and inhibits the attraction of foreign direct investment that could stimulate further growth of the Ukrainian agricultural sector. This, in turn, prevents the Black Sea region from achieving its potential in increasing grain production.

The purpose of this dissertation is to analyze the efficiency of Black Sea grain markets within the framework of price transmission models. The first essay investigates the magnitude and speed of horizontal price transmission of the Black Sea wheat and Ukrainian barley export markets with world grain markets. The second essay investigates short- and long-run price transmission characteristics between world wheat prices and

Ukrainian domestic wheat prices at both producer and consumer levels. The third essay uses the dynamic conditional correlation GARCH model to analyze the dynamics of wheat price volatility in the Ukrainian domestic market.

The notion of “price transmission” refers to the co-movement shown by prices of the same good in different geographic locations. Such analyses have garnered popularity among economists as a tool to study the extent to which rational arbitrage operates. The latter 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 potential profits. This will cease at the point when prices equalize across the different locations, regardless of whether this results in physical trade between different locations. Despite a number of shortcomings that are discussed in the dissertation, price transmission models have been appealing to economists for a number of reasons. Their ability to disentangle short- and long-run price dynamics, and to obtain insightful results from price series only, are among the major ones. Understanding how price shocks are transmitted across and within countries and what affects such transmission serves as a prerequisite for policy analysis. This is yet another reason that price transmission studies are in demand.

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II. BLACK SEA AND WORLD WHEAT MARKET PRICE INTEGRATION ANALYSIS

1. Introduction

During the years leading up to the Russian drought of 2010, the share of Russian and Ukrainian wheat exports in world wheat trade steadily increased. In 2008/09 and 2009/10 these two countries (often referred to as the Black Sea region) together exported an average of 29 million tons of wheat per year (figure 2.1).

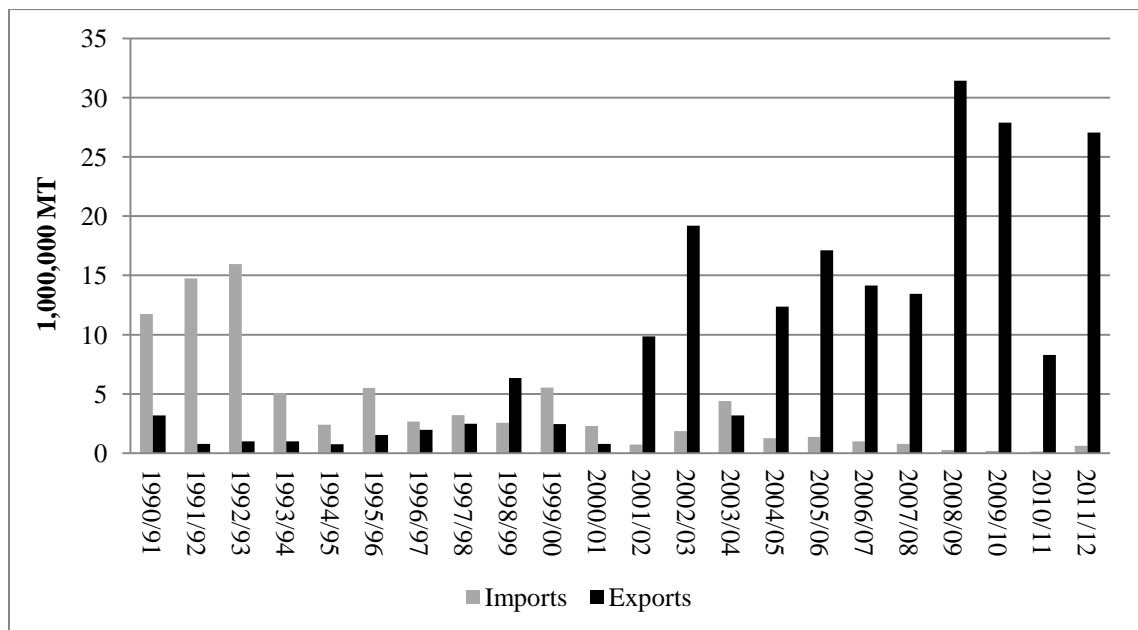


Figure 2.1 Combined Russian and Ukrainian wheat imports and exports

Source: USDA 2013

Such a level of exports accounted for 21.3 percent of the world wheat exports (USDA 2013) and exceeded the level of exports from any other country. After a bad harvest in 2010/11, Ukraine and Russia were back to their position of the largest world exporters at an estimated 25.5 million tons of wheat (USDA 2013) in 2011/12.

The major importers of Ukrainian wheat (2009/10 estimates) were Egypt, Tunisia, Israel, Bangladesh, and Kenya. The primary destinations of the Russian wheat exports were Egypt, Turkey, Iraq, Yemen and Israel with exports of more than 1 million metric tons per country (GTA 2010).

Despite the already large share of Black Sea wheat exports in the world, it could further rise due to an increase in land use and/or yields. An FAO-EBRD (2009) study suggests that up to 6 million hectares of abandoned land in Russia and up to 3 million hectares in Ukraine could be added to wheat production. As to yields, in the last four years both Ukrainian and Russian wheat yields were on average half of those in the EU-27 (USDA 2013), and if improved, can lead to significant growth in wheat production in these countries.

Given the growing significance of Black Sea wheat exports in the international wheat market place, which also coincides with the increasing commodity price levels and volatility around the globe, it is important to learn more about the region's role in wheat market price dynamics. The degree to which prices are transmitted to and from the region might significantly influence not only production incentives in the domestic markets, but can also have an impact on the world market due to the large share of Black Sea wheat in world trade. 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 policy interventions and constructing appropriate models for price analysis.

The objective of this study is to investigate short- and long-run wheat price dynamics between Ukraine and Russia and other major wheat exporters - United States, EU, and Canada. In particular, in this paper we address four major issues. The first

hypothesis we test is that the Black Sea wheat prices are not cointegrated with the wheat markets of the United States, Canada, and EU. There are a number of factors that could prevent prices from convergence to the long-run equilibrium, such as, transaction costs, market power, policy interventions, exchange rates, quality differences, etc. Among the most relevant ones to the Black Sea wheat region are wheat quality heterogeneity, and restrictive export policies. Wheat exported from Russia and Ukraine is of a lower quality than wheat traded by other major exporters. For example, milling quality wheat of Class 4 constitutes the highest share of Russian wheat exports. In the case of Ukraine, feed quality wheat and milling quality wheat of Class 3 account for majority of its exports. With regards to export policies, both countries tend to implement export restrictions in an ad hoc manner. During the financial crisis in 2006-08, the Ukrainian government changed its decisions on wheat export restrictions at least eight times. In 2010, Ukrainian policy makers could not come up with a stable export policy for more than a year, changing the legislation several times. Such an unstable policy environment not only distorts farmers' incentives locally and inhibits the attraction of foreign direct investments, but could also affect cointegration of Black Sea prices with the prices of other exporters. To check for the presence of cointegration we use the Johansen Maximum Likelihood (ML) and residual based tests.

The second question addressed in this paper is whether, for cointegrated pairs of prices, price transmission is symmetric. The threshold autoregressive and momentum threshold autoregressive models are used in this case. Third, with the help of the error correction model we investigate the short run dynamics among the analyzed series that are found to be cointegrated. The results of the above tests will allow us to gain insights

into how fast the Black Sea wheat prices adjust to the changes in the world wheat prices and analyze the possible policy issues that could stem from different levels of price transmission. Fourth, we apply the Bai and Perron (2003) dynamic programming algorithm to check for the presence of structural breaks in the pairs of investigated series, and analyze their possible causes and effects.

Despite their widespread use and increase in complexity, cointegration models have a number of shortcomings that need to be taken into account when interpreting the results. For example, most of these models rely on price data only, and, therefore, a number of assumptions need to be made about their specifications. Most commonly, transaction and transportation costs are assumed to be either equal to zero or set as a fixed proportion of the prices used. They are also assumed to be stationary and serially uncorrelated. Additionally, it is important to differentiate between price cointegration and market integration. Often in literature these two concepts are used as synonymous. However, Barrett (1996) states that “cointegration is neither necessary nor sufficient for market integration”. Instead he differentiates between “market integration” and “competitive market equilibrium”. Market integration is a quantity based indicator that reflects “tradability of products between spatially distinct markets, irrespective of the existence of absence of spatial market equilibrium” (Barrett and Li 2002). On the contrary, competitive market equilibrium or market efficiency is a price-based indicator and holds on the condition of spatial arbitrage. The latter 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, regardless of whether this results in physical trade

between different locations. The models used in this paper utilize price data only, and, thus, serve as analytical tools to evaluate wheat market efficiency in the Black Sea region, rather than its integration with other wheat markets.

Another issue that needs to be accounted for in the interpretation of the results is non homogeneity of different types of wheat. Larue (1991) shows in his research that wheat protein content has a significant influence on wheat prices. As such, quality differences could interfere with price linkages in the international wheat market by making varieties of wheat imperfect substitutes of one another (Mohanty et al. 1999; Shoshray and Lyoid 2003). Nevertheless, wheat prices should still be interrelated to the extent of substitution of different wheat types in their end consumption.

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, little research effort has been dedicated to investigation of Black Sea region wheat price dynamics. To our knowledge only the study by Sagidova (2004) analyzes wheat price dynamics between Ukraine and the United States. Her results suggested that the two series were cointegrated of order one. The short-term relationship, however, was found to be statistically insignificant, which could be explained by the fact that Ukrainian exports were much smaller at that time (up to 2003) compared to the recent ones (see figure 2.1).

Another study by Ghoshray (2008) focused on the analysis of price cointegration between the Ukrainian wheat market and the U.S. corn market. As was mentioned earlier, Ukrainian wheat is feed grade wheat. Therefore, it could be reasonable to assume that the Ukrainian wheat market would be more closely linked to feed grains markets, i.e.

corn and barley, rather than to the food wheat market. However, Ghoshray's findings suggested that there is no cointegration between the given series.

The study by Götz et al. (2013) uses the Markov switching vector error correction model to analyze the effects on the level of price transmission from world prices to farm prices of the export restrictions in 2007-08 that were imposed by Ukrainian and Russian governments. The study shows that while the series remained cointegrated during the whole period that was analyzed, the magnitude of the price transmission to the growers did change in the 2007/08 crop year.

Our paper adds to the above studies by conducting a comprehensive analysis of price transmission levels (both short- and long- run) for Black Sea wheat from 2004 to 2010, and serves a goal to enhance understanding of wheat price dynamics in the region, which has recently grown to be the largest wheat exporting region in the world. Our analysis is used to identify wheat market inefficiencies in the Black Sea wheat market, and serves as a prerequisite for policy recommendations on how to alleviate them. This is also the first study that checks for asymmetry in price transmission in this market. Moreover, our study also contributes to the literature by estimating long-run price transmission elasticities that could be used to better link wheat prices of the Black Sea region to other prices in modeling global wheat market behavior.

The rest of the article is organized as follows: The next sections briefly highlight 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.

2. Methods

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. In this paper, in order to test for the long-run relationship between Russian/Ukrainian and US, Canadian, and EU export 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 co-integrating relationships and on different pairs of the wheat price series.

To be able to test two price series for the co-integrating vector, we first need to confirm the presence of a unit root within each series, which indicates that 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 number of lags was estimated by minimizing the Schwarz Bayesian Criteria (SBC) starting with 12 lags in the initial regressions due to the monthly nature of the 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 KPSS unit root tests (Kwiatkowski et al. 1992) were run as a robustness check of the results obtained from the ADF and PP tests. 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 might be needed (Cheung et al 1994).

Johansen's cointegration test (Johansen 1988) is commonly used to test for the presence of co-integrating vectors. To obtain the test results, we first specify the general VAR(k) model, where k is the number of lags:

$$P_t = A_0 + A_1P_{t-1} + \dots + A_kP_{t-k} + u_t \quad t = 1, \dots, T, \quad (1),$$

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

$$\Delta P_t = \Pi_0 + \Pi P_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \theta_t, \quad (2),$$

where Δ denotes first difference, $\Pi_0 = A_0$, 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 co-integrating 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 co-integrating 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 Russian and US wheat prices:

$$P_t^{RUS} = \beta_0 + \beta_1 P_t^{US} + \varepsilon_t \quad (3),$$

where P_t^{RUS} , P_t^{US} are prices of Russian and US wheat, respectively. β_0 is a constant, β_1 stands for the price transmission elasticity, and ε_t is the error term. Second, we use unit-root tests (ADF, PP, and KPSS) to check if the residuals are stationary. Their stationarity would imply that analyzed series are cointegrated, i.e. they move together in the long-run. If two series are cointegrated, then the OLS estimators in (3) are superconsistent and can be used to characterize the series' behavior. We used both the residual-based and Johansen ML tests for the robustness check. To test for autocorrelation of the residuals the Breusch-Godfrey (Breusch 1979) test was used.

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 setting up a following equation, which is known as the threshold autoregressive (TAR) model:

$$\Delta \bar{\varepsilon}_t = I_t \gamma_1 \bar{\varepsilon}_{t-1} + (1 - I_t) \gamma_2 \bar{\varepsilon}_{t-1} + \varphi_t, \quad (4),$$

where $\Delta \bar{\varepsilon}_t$ is the first difference of the error term from (3), $\bar{\varepsilon}_{t-1}$ is lagged error term from (3) lagged for one time period, γ_1 and γ_2 are the adjustment rates,

$$I_t = \begin{cases} 1 & \text{if } \bar{\varepsilon}_{t-1} \geq \tau \\ 0 & \text{if } \bar{\varepsilon}_{t-1} < \tau \end{cases} \quad (5),$$

and τ is the estimated threshold using Chan's (1993) method.

This method arranges the estimated residual series, $\Delta\bar{\varepsilon}_t$, in ascending order and excludes 15 percent of the smallest and 15 percent of the largest observations. The remaining 70 percent of the values are considered as a possible threshold. For each one of them the equation was estimated based on (4) and (5). A super-consistent estimate, τ , is the one that yields the smallest residual sum of squares.

In some instances instead of searching for the consistent threshold, τ could be set to equal to zero. However, as Tong (1983) mentions, if the price adjustment is asymmetric, the sample mean is not equal to zero, and the use of a super-consistent threshold helps to solve this problem. Equation (4) 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.

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, momentum threshold autoregressive (M-TAR) 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 (5) is modified in the following way:

$$I_t = \begin{cases} 1 & \text{if } \Delta\bar{\varepsilon}_{t-1} \geq \tau \\ 0 & \text{if } \Delta\bar{\varepsilon}_{t-1} < \tau \end{cases} \quad (6),$$

where, as previously, τ is the estimated threshold using Chan's (1993) method. The 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 (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. Therefore, 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 imply that increases in the price of the dependent variable tend to persist, while the decreases are transmitted 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 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 Russia and the U.S.):

$$\Delta P_t^{RUS} = \alpha_0 + \alpha_1 \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \delta_i \Delta P_{t-i}^{RUS} + \sum_{j=1}^n \theta_j \Delta P_{t-j}^{USA} + \mu_t \quad (7),$$

where ΔP_t^{RUS} and ΔP_t^{USA} are vectors of the first differences of log prices for Russia and the US, $\bar{\varepsilon}_{t-1}$ is the lagged residual from (3), μ_t is the error term, and the scalar α_1 represents the short-run adjustment speed of the dependent variable to the long-run steady state (Baffes and Gardner 2003). Its sign is expected to be negative. Equation 7 can be specified with additional lags ($\sum_{i=1}^p \delta_i \Delta P_{t-i}^{RUS}$ and $\sum_{j=1}^n \theta_j \Delta P_{t-j}^{USA}$) 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 (7) since it was found to be statistically insignificant. 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.

Following the procedure used by Ghoshray (2002) we can estimate the number of months (n) it takes Russian and Ukrainian series to adjust back to the equilibrium after the change in US or French prices. The formula to use is the following:

$$n = \frac{\log(1-p)}{\log(1-\alpha_1)} \quad (8),$$

where p is a given proportion of the disequilibrium to be corrected, and α_1 is the short-run adjustment speed coefficient from (7).

3. Data

In our empirical investigation we used monthly wheat FOB prices for Russian Soft Wheat and Ukrainian Feed Wheat (Black Sea ports), Canadian Western Red Spring Wheat (St. Lawrence), U.S. Soft Red Winter Wheat (Gulf ports)¹, and French Soft Wheat (Rouen). The time span of the analysis is from September 2004 till October 2010. We assumed the French wheat price to be representative of the EU wheat price, since it is the largest producer of wheat in the EU (EUROSTAT 2012). All the series are quoted in nominal USD per ton and are expressed in logs. Using prices already converted to the USD and not introducing the exchange rate as a separate repressor is a widely used procedure in spatial price transmission analysis of commodity markets. In those cases when such prices are not expressed in the same currency, research shows that results are generally not altered if the exchange rate is included in a model (Thompson 1999; Bukenya and Walter 2005). We obtained the series from the International Grains Council.

A visual inspection of the graph of the analyzed series in USD per ton suggests that all the series tend to move together over the analyzed period (figure 2.2). The reason for the Canadian prices to be consistently higher than other prices is due to the differences in the quality between soft and hard wheat. We also did not detect a significant trend in any of the series.

¹ Even though the majority of wheat exported by the U.S. is Hard Red Winter wheat, we focused our analysis on wheat types that more comparable in their quality. This defined our choice of the U.S. Soft Red Winter Wheat as a representative U.S. price.

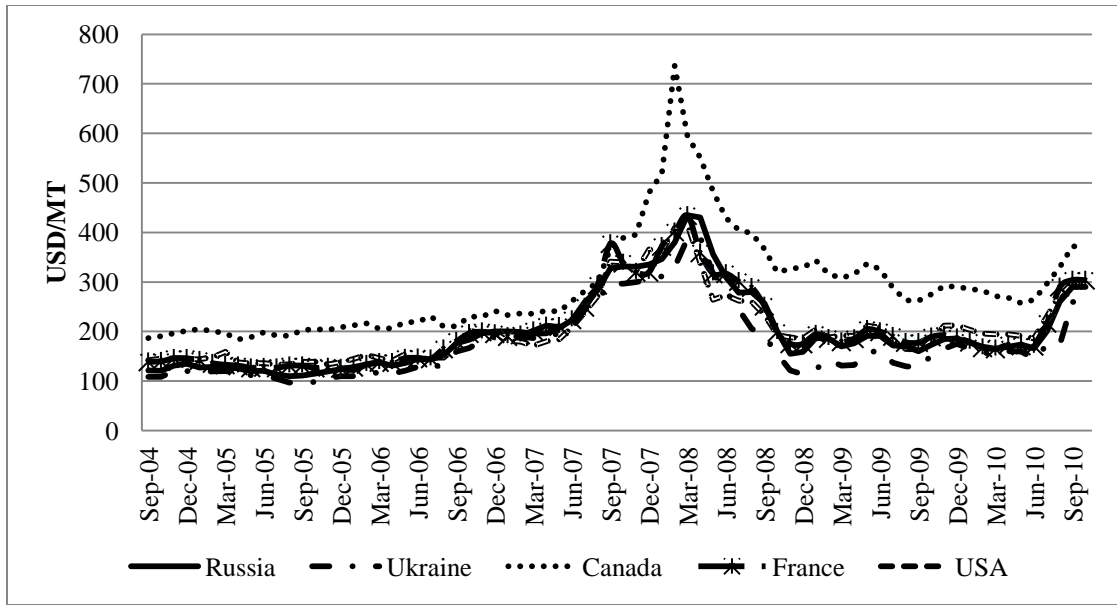


Figure 2.2 Comparison of wheat prices at different export points, \$ per ton

Source: International Grains Council, 2011

The descriptive statistics and characteristics of the analyzed time series is provided in table 2.1. The most volatile price series is the Canadian price. The total number of observations is 74.

Table 2.1 Descriptive statistics and characteristics of the price series (\$ per ton)

Country	Wheat type	Min	Max	Mean	SD	Variance	Skewness	Kurtosis
Ukraine	Feed Wheat	97.0	390.0	174.2	72.5	5268.5	1.3	0.9
Russia	Soft Wheat	110.0	435.0	198.6	79.1	6266.0	1.4	1.3
Canada	Western Red Spring	184.2	736.2	290.8	105.3	11092.1	1.8	4.2
USA	Soft Red Winter	133.2	415.3	204.6	69.3	4812.8	1.4	1.2
France	Standard Grade Wheat	126.1	434.4	205.1	76.3	5830.8	1.2	0.6

4. 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 (Table 2.2). Thus, the tests were re-run on the series after they were differenced in log levels.

Table 2.2 Results of the unit root tests in levels^a

	# of lags	ADF		PP		KPSS	
		w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
Ukraine	1	-1.61	-1.77	-1.26	-1.42	0.58**	0.28**
Russia	4	-1.78	-2.16	-1.37	-1.60	0.75**	0.31**
France	3	-1.30	-1.59	-1.27	-1.57	0.74**	0.29**
US	1	-1.74	-2.11	-1.38	-1.72	0.80**	0.25**
Canada	2	-1.70	-2.09	-1.46	-1.66	0.95**	0.27**

^aAsterisks 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.

The results provided in table 2.3 show that all the differenced series are stationary. This leads to the conclusion that the price series of Russia, Ukraine, US, Canada and EU are I(1). The tests were run for the cases when trend is present, and when it is absent. The results show that including the trend does not affect the outcome.

Table 2.3 Results of the unit root tests using first difference^b

	# of lags	ADF		PP		KPSS	
		w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
Ukraine	2	-3.91**	-3.86**	-4.77**	-4.73**	0.10	0.11
Russia	2	-3.64**	-3.60**	-6.06**	-6.02**	0.11	0.11
France	2	-4.36**	-4.32**	-5.15**	-5.12**	0.10	0.11
US	1	-5.31**	-5.27**	-6.06**	-6.02**	0.08	0.08
Canada	1	-4.29**	-4.26**	-6.74**	-6.70**	0.09	0.08

^bAsterisks 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 2.4 suggest that we reject the null hypothesis of no cointegration, i.e. $r = 0$, where r is the number of co-integrating relationships. We also reject H_0 that there is only one co-integrating relationship ($r = 1$). Since, however, we fail to reject the null hypotheses that $r = 2, 3$ or 4 , we conclude that there are two or more distinct long-run relationships among the five series.

Table 2.4 Cointegration rank test using trace^c

Ho(Rank=r)	H1(Rank>r)	Trace	5% CV
0	0	112.08**	75.74
1	1	63.47**	53.42
2	2	27.87	34.80
3	3	12.26	19.99
4	4	5.89	9.13

^cAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

In order to find out which pairs of series are cointegrated, both Johansen's ML and Engle-Granger cointegration tests were run on the pairs of series. The Engle-Granger tests suggest that the cointegrated pairs of series are Russia-France, Russia-USA², and Ukraine-France (table 2.5). In the Ukrainian case, the ADF test results do not support the presence of cointegration between Ukraine and France. However, since two of the three (PP, KPSS and Johansen ML) tests show evidence of cointegration being present, we conclude that two series are cointegrated.

² We also ran the cointegration tests for the U.S. HRW wheat, however, no statistically significant long-run relationship was found between these prices and those of Russia and Ukraine

Table 2.5 Engle- Granger cointegration tests for the wheat price series of interest^d

Pair of series	# of lags	ADF	PP	KPSS
Russia-France	2	-5.32**	-5.24**	0.11
Russia-Canada	1	-2.30	-2.38	0.24
Russia-USA	1	-3.79**	-3.81**	0.23
Ukraine-France	3	-2.33	-3.64*	0.28
Ukraine-Russia	3	-2.00	-2.22	0.46*
Ukraine-USA	2	-2.91	-3.24*	0.37*
Ukraine-Canada	1	-1.90	-1.99	0.26

^dAsterisks 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 Russian-French and Ukrainian-French pairs of wheat prices, however, not the Russian-US relationship (table 2.6).

Table 2.6 Johansen ML Pairwise cointegration tests for the wheat price series of interest^e

Pairs of series	Ho(H1)	Trace	5%CV
Russia-France	r=0((r>0)	25.98**	19.99
	r=1(r>1)	6.69	9.13
Russia-Canada	r=0((r>0)	13.23	19.99
	r=1(r>1)	5.12	9.13
Russia-USA	r=0((r>0)	15.47	19.99
	r=1(r>1)	3.75	9.13
Ukraine-France	r=0((r>0)	24.66**	19.99
	r=1(r>1)	5.28	9.13
Ukraine - Russia	r=0((r>0)	14.60	19.99
	r=1(r>1)	2.69	9.13
Ukraine-USA	r=0((r>0)	15.51	19.99
	r=1(r>1)	4.70	9.13
Ukraine - Canada	r=0((r>0)	12.48	19.99
	r=1(r>1)	4.56	9.13

^eAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

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 Russian and US wheat series are also cointegrated. Lack of cointegration between Russia-Canada and Ukraine-Canada pairs of prices is most likely attributed to the significant difference in quality of Canadian hard and Black Sea soft wheat.

Since Russia-France, Russia-US and Ukraine-France series are cointegrated, the results of the regressions that analyze the relationships between them are consistent (see equation 3). Thus, β_1 can be considered as the long-run price transmission elasticity. The results show that the long-run price elasticity is 1.04 between Russian and French wheat series, 1.16 between Russian and US wheat series, and 1.05 between Ukrainian and French wheat series. The price transmission elasticity indicates the percentage change in the price of one country in response to a one-percent change in another country's price. It is directly related to trade liberalization (Listorti 2008; Thompson 1999), since higher levels of trade liberalization contribute to greater price transmission elasticities. Of course, such an interpretation needs to be introduced with caution. As was mentioned in the introduction, for those price transmission models that only use price data, an assumption is made that transportation or transactions costs are proportional to the price of good (if prices are expressed in logs). This is rather strong assumption that will almost surely not hold in the real world.

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 three pairs of prices, we reject the null of no cointegration ($H_0: \gamma_1 = \gamma_2 = \mathbf{0}$) based on the Φ -statistic estimates, and confirm our previous results that the Russian wheat price series are cointegrated with the US and French wheat prices, and Ukrainian prices are cointegrated

with French ones (table 2.7). The F-statistics for both TAR and M-TAR model suggest that we cannot reject the null of the symmetric price transmission (tables 2.7 and 2.8). Therefore, we conclude that the price transmission between the three pairs of prices is symmetric.

Table 2.7 TAR model parameter estimates^f

Pair of prices	Russia - France	Russia-USA	Ukraine –France
Variable	Parameter estimate	Parameter estimate	Parameter estimate
γ_1	-0.80 (-5.31)**	-0.26 (-2.05)*	-0.22 (-1.18)
γ_2	-0.74 (-4.76)**	-0.36 (-2.30)*	-0.38 (-2.75)**
$H_0: \gamma_1 = \gamma_2 = 0(\Phi)$	22.08**	4.38**	6.82**
$H_0: \gamma_1 = \gamma_2 (F)$	0.11[0.74]	0.24 [0.62]	0.33 [0.57]
τ	-0.019	-0.04	0.016

^fAsterisks 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 2.8 M-TAR model parameter estimates^g

	Russia - France	Russia-USA	Ukraine –France
Variable	Parameter estimate	Parameter estimate	Parameter estimate
γ_1	-0.50 (-2.10)**	-0.24 (-2.19)**	-0.29 (-2.26)**
γ_2	-0.79 (-3.73)**	-0.46 (-3.33)**	-0.43 (-2.69)**
$H_0: \gamma_1 = \gamma_2 = 0(\Phi)$	6.94**	7.93**	5.21**
$H_0: \gamma_1 = \gamma_2 (F)$	1.89 [0.17]	1.50 [0.23]	0.57 [0.45]
τ	-0.002	-0.023	-0.069

^gAsterisks 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.

The absence of asymmetry in 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 results show (table 2.9) that the error-correction terms for Russian prices are negatively significant. This suggests that U.S. and French prices both adjust to the long-run equilibrium of Russian prices. Using equation (9) we estimate that it takes French wheat price about 2.6 months to correct 90% of

disequilibrium after the change in Russian price. For the USA the adjustment duration is much slower and is equal to 6.4 months. On the contrary, the error correction terms are not significant for Russian prices both in Russia-France and Russia-USA equations. This implies that Russian prices do not adjust to correct long-run disequilibria, and serves as an indication that Russia is a price leader in its relationship with French and U.S. series. The fact that Russia might serve as a price leader in the analyzed relationships is not surprising due to the large shares of Russian exports in the analyzed period. Except for the 2004/05 marketing year, Russian annual wheat exports exceeded French wheat exports on average by approximately 2 million tons. In case of the U.S., during the entire analyzed period, the amount of Russian wheat exports exceeded that of the U.S. soft wheat exports by an average of 8 million tons (USDA 2013, EUROSTAT 2012).

Table 2. 9 Error-correction model parameter estimates^h

Dependent variable	Independent variable	# of lags	Speed of adjustment, α_1	LM test	F _{test}
Russia	France	2;2	-0.18 (-0.82)	1.87 [0.76]	12.78**
France	Russia	1;1	-0.59 (-2.82)**	1.73 [0.78]	10.08**
Russia	USA	1;1	-0.09 (-0.98)	3.28 [0.51]	14.78**
USA	Russia	1;1	-0.30 (-2.11)**	1.66 [0.79]	4.03**
Ukraine	France	2;2	-0.20 (-1.99)*	0.92 [0.63]	11.38**
France	Ukraine	1;1	-0.25(-2.63)**	2.76 [0.25]	9.13**

^hAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

In case of the relationship between Ukrainian and French wheat prices, approximately 20-25 percent of adjustment occurs in one month regardless of which country initiates the price change. Thus, we can conclude that neither country can be regarded as a price leader. Each price adjusts to the equilibrium after the change in another price in approximately 10 months.

The difference in the time of adjustment between Russian and Ukrainian prices to the change in the French price is surprising. Closer visual examination (figure 2.2),

however, shows that during the price spike of 07/08, the difference between Ukrainian and French prices significantly increased, and such increased deviation persisted till the beginning of 2010. At the same time Russian wheat prices have continued to closely follow the French series. Such a behavior in the Ukrainian wheat price series could have accounted for the lower rate of the adjustment in the ECM, and suggest a possible structural break.

To check for the presence of such a break for the Ukraine-French pair of prices³, we employed the Bai and Perron (2003) dynamic programming algorithm. Based on the BIC and RSS criteria, three breaks were found. However, due to the relatively small number of observations we restricted our analysis to only one break, which fell on June 2008 with a 95 percent confidence interval for the estimated break points stretching from June 2007 till September 2008. As shown in figure 2.3, this break coincides with the commodity price bubble and with more frequent interventions by the Ukrainian government in the domestic grain trade policy that could possibly be attributed to the increase in the price difference between Ukraine and France.

Regardless, we can still analyze the effects of the break on the long-run price transmission elasticity between two series before and after the break. For this purpose, we re-run cointegration tests on two periods - from September 2004 till May 2008 (pre-break period), and from June 2008 till October 2010 (post-break period).

³ We also applied same procedure to the pairs of prices that were found to be not cointegrated (i.e. Ukraine-USA, Ukraine-Canada, or Russia-Canada pairs) to check if the presence of the structural break could have improved cointegration. However, the inclusion of the structural breaks in these long-run relationships did not show statistically significant presence of cointegration among the series.

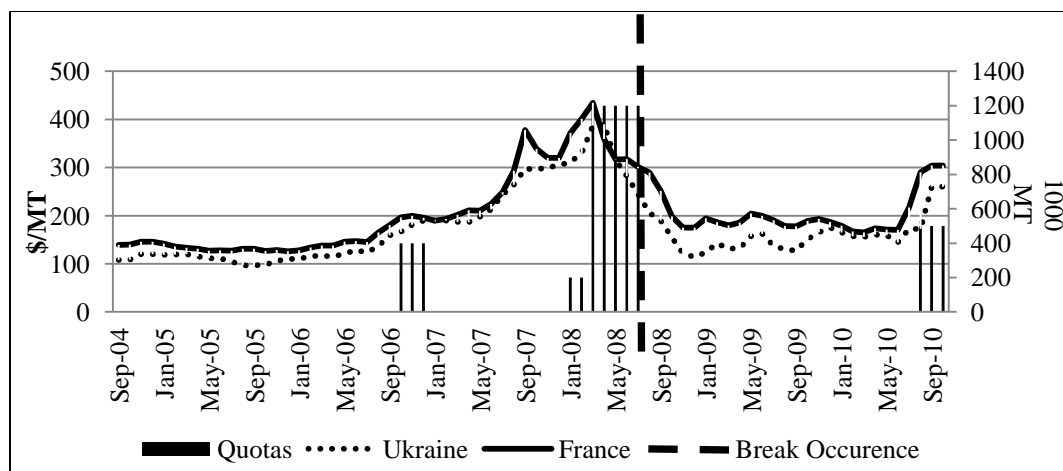


Figure 2.3 Chronology and size of the Ukrainian wheat export quotas and an occurrence of the structural break ⁴

Source: International Grains Council, 2011; FAO-EBRD, 2010.

The results, summarized in table 2.10, suggest that there is lack of price cointegration between two series in the post-break period. However, the series in the pre-break period are still cointegrated ⁵. The long run price transmission elasticity in this period is equal to 1.06, which is similar to the result obtained for the entire analyzed period.

Table 2.10 Engle-Granger cointegration test results before and after the break in the Ukraine-France long-run relationship

	# of lags	ADF	PP	KPSS
Pre-break period	4	-2.61	-3.14*	0.17
Post-break period	2	-1.75	-2.74	0.35*

¹Asterisks 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.

⁴ From December 14, 2006 till February 15, 2007, and then from July 1, 2007 till January 1, 2008 the amount of wheat export quota was equal to 3,000 tons

⁵ Johansen maximum likelihood test results supported same conclusion. Therefore, even though ADF test results were not statistically significant in the pre-break period, based on the results of PP, KPSS and Johansen ML tests, we suggest that series are cointegrated in this period.

5. Conclusions and implications

The study examines the long- and short-run dynamics between Black Sea and the US, Canadian, and EU wheat prices using monthly FOB data from July 2004 to October 2010. The results have important implications for evaluating the performance of the Black Sea grain market that has recently emerged as a major world grain export region. Speed and magnitude of price transmission between countries show how well the price signals are transmitted to and from different markets and define the allocation of resources for various economic activities. Lack or slow speed of price transmission might result in potential losses to different players in the Black Sea wheat market, and signify the presence of the market inefficiencies that need to be corrected.

Results of our analysis show that Russian wheat prices are cointegrated with French and US wheat prices, but not with the Canadian ones, which is most likely attributed to the difference in quality and final uses between Russian and Canadian wheat. Ukrainian wheat prices were found to be cointegrated with French wheat price series as well but not with the US or Canadian ones. The estimated long-run price transmission elasticities are equal to 1.04 between Russian and French series, 1.16 between Russian and US series, and 1.05 between Ukrainian and French series.

Furthermore, the results of the ECM model suggest that Russia could possibly serve as a price leader in its relationship with the U.S. and French soft wheat. One possible explanation for such results is that during most analyzed years Russian wheat export amounts exceeded those of France in general as well as in two major regions (Near East Asia and North Africa) where Russia and France directly compete for markets. Moreover, over 30 percent of Russia's total exports (2009-2010 estimates) are

shipped to Egypt which is one of the largest importers of wheat and is an important country for wheat price formation in Atlantic trade. Speed of adjustment towards the equilibrium after the change in Russian wheat price is equal to 2.6 months for France and 6.4 months for the U.S.

With regards to Ukraine, based on our results we can conclude that the commodity price spike of 2007-2008, which also caused the Ukrainian government to heavily intervene in the wheat exports, coincided with a structural change for the price relationship between Ukrainian and French prices. This, in turn, resulted, in a lack of cointegration between these two series in the post-change period.

The results of the TAR and M-TAR models suggested that shocks to the EU and/or US price are transmitted to the Russian and Ukrainian wheat prices symmetrically. Such results were not surprising, since international wheat markets are rather dynamic without any single country being able to exhibit distinct market power and control wheat prices in the world markets. However, investigation of symmetry in price transmission should be further extended to the analysis of the domestic wheat markets in Ukraine and Russia. More specifically, it would be interesting to see how well the wheat export price is transmitted to domestic consumers and farmers.

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III. SHORT- AND LONG- RUN RELATIONSHIPS BETWEEN UKRAINIAN BARLEY AND WORLD FEED GRAIN EXPORT PRICES

1. Introduction

In the early 1990s, the European Union (EU) was a leading barley exporter, accounting for about 44 percent of the total export barley market share (USDA 2013). Two other major exporters, Australia and Canada, had to compete with the EU’s logistical advantages and farm support programs that guaranteed high prices for EU producers. However, just ten years later, the dynamics in the world barley market has changed dramatically. Significant decreases in agricultural support to EU grain producers and the emergence of Black Sea grain exporters resulted in a change of leadership in the global barley market. Between 2004/05 and 2012/13 its new leaders were Ukraine and Australia with each one of them exporting on average 25 percent of the total world barley annually (figure 3.1).

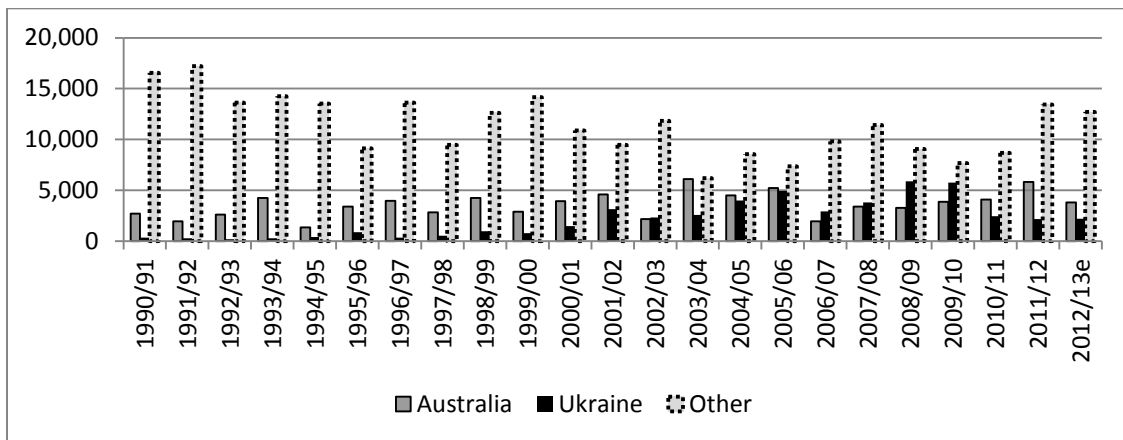


Figure 3.1 World barley exports between 1990/91 and 2012/13, 1000 MT

Source: USDA, January 2013

Given this new state of affairs in the international barley market, the objective of this paper is to report on the role of one of its leading barley exporters, Ukraine, in the international feed grain market price dynamics. Specifically, we employ price transmission techniques to analyze how efficiently the barley prices are transmitted between Ukraine (feed barley) and other major feed grain exporters, such as Australia (feed barley), European Union (feed barley), Canada (malting barley) and United States (corn⁶) in both short- and long-run.

The first hypothesis to be tested is that Ukrainian barley prices are not cointegrated with those of Australian, EU, Canadian or U.S. feed grains⁷. There are a number of factors that could prevent prices from convergence to the long-run equilibrium, such as, transaction costs, market power, policy interventions, exchange rates, quality differences, imperfect flows of information, etc. To check for the presence of cointegration we use the Engle-Granger residual based and Johansen ML tests.

The second question addressed in this paper is whether, for cointegrated pairs of prices, price transmission is symmetric. The threshold autoregressive and momentum threshold autoregressive models are used in this case. Different grain market players have periodically raised concerns over the competitiveness of international grain markets, which could be attributed to the fact that a small number of international grain exporters handle the majority of grain exports (Patterson and Abbott 1994; Bessler et al. 2002). The results of the asymmetric price transmission tests in this paper would shed the light on world barley market competitiveness.

⁶ We assume that the Ukrainian barley prices might be cointegrated with the U.S. corn prices due to these two feed grains being substitutes in animal feed production. Therefore, we add corn series to our analysis.

⁷ For the convenience, from here forward we use “feed grains” to denote Australian, EU and Canadian barley as well as U.S. corn.

Finally, with the help of the error correction model we investigate the short run dynamics among the analyzed series that are found to be cointegrated. The results of the above tests will allow us to gain insights into how fast the analyzed series return to the equilibrium after a shock.

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, which is especially relevant for Ukraine, a country whose government frequently uses export restrictions in its grain markets. It should be kept in mind that the magnitude of price transmission parameters obtained from the models used in this analysis summarize the overall effects of all the factors that might affect prices in different markets. Further research is needed to separate the role of these different factors contributing to the various degrees of cointegration. Nevertheless, the obtained parameters and their significance levels provide useful insights about the extent to which different markets share the same price shocks (Conforti 2004).

The choice of the country in itself is important. Despite the already large share of Ukrainian barley exports in the world, it could grow further due to an increase in land use and/or yields. An FAO-EBRD (2009a) study suggests that up to 3 million ha could be added to crop production in Ukraine. As to yields, in the last four years Ukrainian barley yields were on average half of those in the EU-27 (USDA 2013). Increasing grain production in the country would be important not only for the Ukrainian economy, but would also be a valuable addition to the world grain supplies in times of rising food and feed demands around the globe. Focusing our analytical effort on investigating the potential constraints to the Ukrainian barley market efficiency is important as well, since

once identified and corrected, these inefficiencies could further lead to increased production in the country. While analysis of the Ukrainian wheat market efficiency has received some attention from the researchers in recent years (Goychuk and Meyers 2013, Gotz et al. 2013, Goychuk 2013), to our knowledge no studies has been performed on the efficiency of the Ukrainian barley market.

Despite their widespread use and increase in complexity, cointegration models have a number of shortcomings that need to be taken into account when interpreting the results. For example, most of these models rely on price data only, and, therefore, a number of assumptions need to be made about their specifications. Most commonly, transaction costs are assumed to be either equal to zero or set as a fixed proportion of the prices used. They are also assumed to be stationary and serially uncorrelated.

When it comes to the cointegration analysis, one needs to differentiate between price cointegration and market integration. Often in literature these two concepts are used as synonymous. However, Barrett (1996) states that “cointegration is neither necessary nor sufficient for market integration”. Instead he differentiates between “market integration” and “competitive market equilibrium”. Market integration is a quantity based indicator that reflects “tradability of products between spatially distinct markets, irrespective of the existence of absence of spatial market equilibrium” (Barrett and Li 2002). On the contrary, competitive market equilibrium or market efficiency is a price-based indicator and holds on the condition of spatial arbitrage. The latter 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, regardless of whether this results

in physical trade between different locations. The models used in this paper utilize price data only, and, thus, serve as an analytical tool to evaluate efficiency in the global barley markets, rather than their integration.

The rest of the article is organized as follows: The next sections briefly highlight the econometric methods that are 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.

2. Econometric methods

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. In this paper, in order to test for the long-run relationship between Ukrainian barley and US, Canadian, Australian and EU feed grain 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 co-integrating relationships and on different pairs of the barley price series.

To be able to test two price series for the co-integrating 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 number of lags was estimated by minimizing the Schwarz Bayesian Criteria (SBC) starting with 12 lags in the initial regressions due to the monthly nature of the 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 KPSS unit root tests (Kwiatkowski et al. 1992) were run as a robustness check of the results obtained from the ADF and PP tests. 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).

Johansen's cointegration test (Johansen 1988) is commonly used to test for the presence of co-integrating vectors. To obtain the test results, we first specify the general VAR(k) model, where k is the number of lags:

$$P_t = A_0 + A_1 P_{t-1} + \dots + A_k P_{t-k} + u_t, \quad t = 1, \dots, T, \quad (1),$$

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

$$\Delta P_t = \Pi_0 + \Pi P_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \theta_t, \quad (2),$$

where Δ denotes first difference, $\Pi_0 = A_0$, 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 co-integrating 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 co-integrating 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:

$$P_t^{UKR} = \beta_0 + \beta_1 P_t^{AUS} + \varepsilon_t \quad (3),$$

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 use unit-root tests (ADF, PP, and KPSS) to check if the residuals are stationary. Their stationarity would imply that analyzed series are cointegrated, i.e. they move together in the long-run. If two series are cointegrated, then the OLS estimators in (3) are superconsistent and can be used to characterize the series' behavior. We used both the residual-based and Johansen ML tests for the robustness check. To test for autocorrelation of the residuals the Breusch-Godfrey (Breusch 1979) test was used.

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 setting up the following equation, which is known as the threshold autoregressive (TAR) model:

$$\Delta \bar{\varepsilon}_t = I_t \gamma_1 \bar{\varepsilon}_{t-1} + (1 - I_t) \gamma_2 \bar{\varepsilon}_{t-1} + \varphi_t, \quad (4),$$

Where $\Delta \bar{\varepsilon}_t$ is the first difference of the error term from (3), $\bar{\varepsilon}_{t-1}$ is the lagged error term from (3) lagged for one time period, γ_1 and γ_2 are the adjustment rates,

$$I_t = \begin{cases} 1 & \text{if } \bar{\varepsilon}_{t-1} \geq 0 \\ 0 & \text{if } \bar{\varepsilon}_{t-1} < 0 \end{cases} \quad (5).$$

Alternatively, it is also possible that the process of adjustment to the long run equilibrium may depend on the change in $\bar{\varepsilon}_{t-1}$ instead of the level of $\bar{\varepsilon}_{t-1}$ and, therefore, exhibiting more momentum in one direction than another. For such cases, Enders and Siklos (2001) suggested an alternative momentum threshold autoregressive (M-TAR) model in which the Heaviside indicator function is modified in the following way

$$I_t = \begin{cases} 1 & \text{if } \Delta \bar{\varepsilon}_{t-1} \geq 0 \\ 0 & \text{if } \Delta \bar{\varepsilon}_{t-1} < 0 \end{cases} \quad (6)$$

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 (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. Therefore, 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.

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):

$$\Delta P_t^{UKR} = \alpha_0 + \alpha_1 \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \delta_i \Delta P_{t-i}^{UKR} + \sum_{j=1}^n \theta_j \Delta P_{t-j}^{AUS} + \mu_t \quad (7),$$

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 (3), μ_t is the error term, and the scalar α_1 represents the short-run adjustment speed of the dependent variable to the long-run steady state (Baffes and Gardner 2003). Its sign is expected to be negative. Equation 7 can be specified with additional lags ($\sum_{i=1}^p \delta_i \Delta P_{t-i}^{UKR}$ and $\sum_{j=1}^n \theta_j \Delta P_{t-j}^{AUS}$) 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 (7) since it was found to be statistically insignificant. 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.

Following the procedure used by Ghoshray (2002) we can estimate the number of months (n) it takes the analyzed series to adjust back to the equilibrium after a price shock. The formula to use is the following:

$$n = \frac{\log(1-p)}{\log(1-\alpha_1)} \quad (8),$$

where ρ is a given proportion of the disequilibrium to be corrected, and α_1 is the short-run adjustment speed coefficient from (7).

3. 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 to 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 2012). 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.

All the series are quoted in nominal USD per ton and are expressed in logs. Using prices already converted to the USD and not introducing the exchange rate as a separate regressor is a widely used procedure in spatial price transmission analysis of commodity markets. In those cases when such prices are not expressed in the same currency, research shows that results are generally not altered if the exchange rate is included in a model (Thompson 1999; Bukenya and Walter 2005). The series were obtained from the International Grains Council and HGCA.

According to the International Grains Council (2012), in 2009, more than 80% of total Ukrainian barley was exported to Near East Asia (Saudi Arabia, Jordan and Iran). At his time both Australia and France were also exporters to the same region with 32 and 28 percent shares of their total barley exports, respectively. Most of the Australian malting

barley was, however, shipped to China and Japan. As to France, the top destinations for its barley (both feed and malting) included Near East Asia, China and North Africa.

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 3.2), 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 coincides with the barley series.

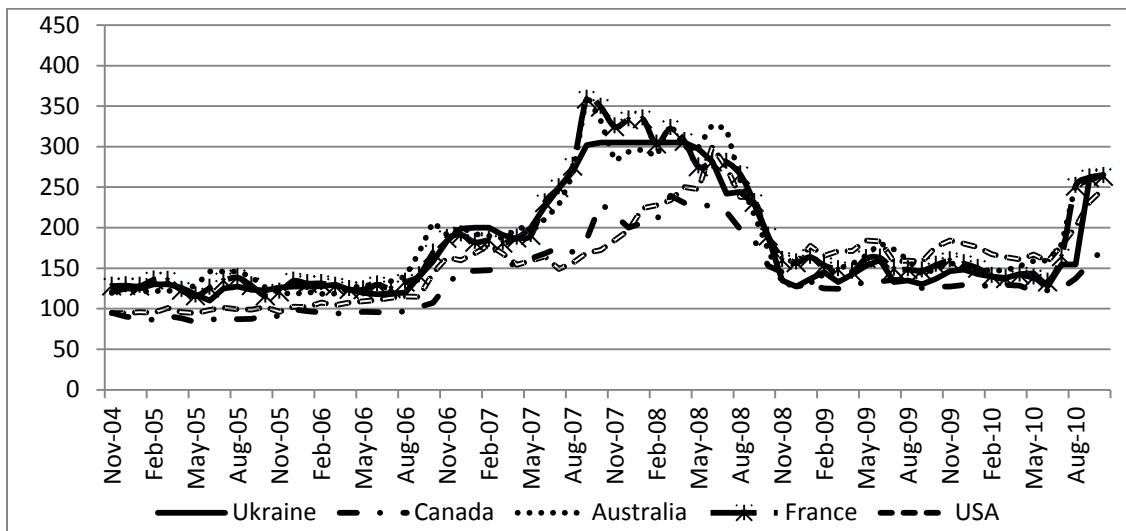


Figure 3.2 Comparison of barley export prices at different export points, \$ per ton

Source: International Grain Council, HGCA 2012

4. 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 (table 3.1). Thus, the tests were re-run on the series after they were differenced in log levels.

Table 3.1 Results of the unit root tests in levels^a

	# of lags	ADF		PP		KPSS	
		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**

^aAsterisks 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 3.2 Results of the unit root tests using first difference^b

	# of lags	ADF		PP		KPSS	
		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

^bAsterisks 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.

The results provided in table 3.2 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 the border line of being insignificant. And since all other

unit root tests' results do not differentiate between the presence and absence of a trend, we conclude that trend inclusion does not affect the outcome.

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 3.3 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 3.3 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

^cAsterisks 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 series⁸. Engle-Granger tests suggest that the cointegrated pairs of series are Ukraine-Australia, Ukraine-France,

⁸ Since the focus of this study is on the Ukrainian barley market, we will not discuss in much detail relationships among other barley exporters' prices.

however, no statistically significant long-run relationship was found for the Ukraine-USA or Ukraine-Canada pairs of prices⁹ (table 3.4).

Table 3.4 Engle- Granger cointegration tests for barley/corn price series^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**

^dAsterisks 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 and Ukraine-France pairs of barley prices (table 3.5).

Table 3.5 Johansen ML Pairwise cointegration tests for barley/corn price series^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

^eAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

Given that the Ukraine-Australia and Ukraine-France pairs are co-integrated, the results of the regressions that analyze their long-run relationships are consistent (see equation 3). Thus, β_1 can be considered as the long-run price transmission elasticity and is equal to 0.71 for Ukrainian and French pair of prices and 0.59 between Australian and Ukrainian barley series. The price transmission elasticity indicates the percentage change

⁹ Other two pairs of prices that were found to be cointegrated include Australia-Canada and Australia-France. Johansen ML test results, however, did not support cointegration of the Australia-Canada pair of prices.

in the price of one country in response to a one-percent change in another country's price. It is directly related to trade liberalization (Listorti 2008; Thompson 1999), since higher levels of trade liberalization contribute to greater price transmission elasticities that are closer to 1. Of course, such an interpretation needs to be introduced with caution. As was mentioned in the introduction, for those price transmission models that only use price data, an assumption is made that transportation or transactions costs are proportional to the price of good (if prices are expressed in logs). This is rather strong assumption that will almost surely not hold in the real world.

The next step is to proceed with testing for asymmetric price transmission in the pairs of prices that were found to be cointegrated. Both TAR and M-TAR models were run with different lag lengths, with the optimal number of lags being selected based on the AIC criterion. In the cases of all four pairs of prices, we reject the null of no cointegration ($H_0: \gamma_1 = \gamma_2 = \mathbf{0}$), and confirm our previous results of long-run relationships between Ukraine-France and Ukraine-Australia pairs of prices (table 3.6).

Table 3.6 TAR and M-TAR model parameter estimates^f

Variable	TAR model		M-TAR model	
	France-Ukraine	Australia-Ukraine	France-Ukraine	Australia-Ukraine
	Parameter estimate	Parameter estimate	Parameter estimate	Parameter estimate
γ_1	-0.93 (-4.25)**	-0.54 (-2.16)**	-1.15 (-6.52)**	-0.89 (-3.26)**
γ_2	-1.31 (-6.69)**	-0.97 (-4.70)**	-0.78 (-3.88)**	-0.39 (-1.84)**
$H_0: \gamma_1 = \gamma_2 = 0(\Phi)$	25.43**	11.37**	28.54**	6.39**
$H_0: \gamma_1 = \gamma_2 (F)$	2.35 [0.13]	2.55 [0.12]	2.00 [0.16]	2.41 [0.13]

^fAsterisks 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.

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. Therefore, we conclude that the price transmission between the four pairs of prices is symmetric.

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 (table 3.7) 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 prices¹⁰.

Table 3.7 Error-correction model parameter estimates^h

Dependent variable	Independent variable	# 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**

^hAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

Thus, 90 percent of the full adjustment takes about 6 months in the case of Ukraine-Australia pair of prices. In case of France, adjustment is slightly faster – 90 percent happens in 5 months. If compared to the literature on short-run price adjustment in grain markets (Ghoshray 2007, Dawson et al. 2006), a 5-6 month adjustment to the equilibrium could be considered as an average magnitude. The fact that Ukraine serves as a price leader in feed barley market is not very surprising. More than 94 percent of all

¹⁰ Given the opposite timing of grain production seasons in Australia and Ukraine, we introduced monthly dummies in the ECM estimation, however, they were found to be insignificant.

barley produced in Ukraine is feed barley, making malting barley exports almost inexistent. On the contrary, in both France and Australia the barley export is split between feed and malting barley at about 60 to 40 ratio (FAO-EBRD 2009b).

5. Conclusions and future developments

We investigated the long- and short-run dynamics between Ukrainian and Australian, Canadian, EU and US feed grain prices using monthly FOB data November 2004 till October 2010. The results suggest that Ukrainian barley prices are cointegrated with Australian and French prices. The estimated long-run barley price transmission elasticity is 0.71 between Ukrainian and French (a representative country of the EU) barley prices, and 0.59 between Australian and Ukrainian 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. This is an average rate of adjustment compared to results of other studies.

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 in the given period. The results of the TAR and M-TAR models also show that shocks between French, Australian and Ukrainian barley prices are transmitted symmetrically.

Interestingly enough, however, the post 2010 developments in the world barley market might suggest further change in its dynamics. Since 2010 the share of the

Ukrainian barley exports has been declining. In 2011/12 and 2012/13 Ukraine occupied only the 5th place among the top barley exporters, while the world saw a sudden emergence of a new leader – Argentina. In just two years between 2009 and 2011, Argentinean exports increased twofold from about 600 thousand tons to 1.2 million tons. In the same time period, exports to Saudi Arabia increased from 4 thousand tons to 300 thousand tons, making Argentina a primary competitor to Ukrainian exporters in the region.

So, what are the factors that are affecting Ukrainian performance in the post 2010 period? Anecdotal evidence suggests that apart from unfavorable weather conditions in both 2010/11 and 2012/13, more restrictive export policies with regards to barley than for the other grains and overall uncertainty about future export policies in the country were among the major causes causing farmers to switch to corn production (USDA FAS 2011, 2013). Additional research is, however, needed to empirically investigate the role of export restrictions on price dynamics in the Ukrainian barley markets. The major conclusion that this study offers is that under the right set of the production incentives, Ukraine has a great potential in barley production and exports.

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IV. ANALYSIS OF THE ASYMMETRIC PRICE TRANSMISSION IN THE UKRAINIAN WHEAT SUPPLY CHAIN

1. Introduction

Analysis of price transmission along the supply chain has become a frequently used method by economists to study the efficiency of markets and identify price response rigidity in the marketing system. The purpose of such analysis is the determination of the price linkages across different marketing levels (i.e. exporter-producer-consumer), as well as the extent and the speed at which a price change that occurs at one stage is also reflected at other stages. If price transmission is not perfect, this implies that the markets may convey inaccurate information to producers and consumers, thus leading to misallocation of resources in the economy.

This paper aims to identify potential market and policy failures in the Ukrainian soft wheat supply chain that might inhibit its efficiency. Ukraine constitutes a particularly interesting case. It has recently emerged into becoming one of the largest world grain exporters. In the 2012/13 marketing year, according to USDA (2013) Ukraine was among the top ten global suppliers of corn (13.5 mln tons), wheat (7 mln. tons), and barley (2.2 mln tons). The expansion into the world markets, however, did not come with a similar openness in the approach of the Ukrainian government to policy implementation. Until recently it has been characterized by frequent and rather ad hoc policy interventions into the Ukrainian grain exports. For example, during the financial crisis in 2006-08, the Ukrainian government changed its decisions on wheat export restrictions at least eight

times. In 2010, Ukrainian policy makers could not come up with a stable wheat export policy for more than a year, again changing the legislation several times. Such an unstable policy environment both distorts farmers' incentives locally and restrains the attraction of foreign direct investments. Also, as previous research shows (Goychuk and Meyers 2013), it affected cointegration of Black Sea prices with the prices of other exporters. However, policy makers justify the abovementioned controls as the ones that protect Ukrainian consumers from the high and volatile world grain prices.

More specifically, the goal of this paper is to investigate the price transmission between the world wheat prices and Ukrainian domestic wheat prices at both producer (farm) and consumer (flour) levels. First, we analyze the transmission of the world prices into the Ukrainian domestic wheat market under different policy restrictions and identify potential losers in the marketing chain. The hypothesis is that there is a lack of transmission between these two series that could be caused by a number of factors, such as, transaction costs, market power, policy interventions, exchange rates, quality differences, etc.

Second, we investigate the price transmission patterns from wheat producers to flour millers. Two hypotheses are to be tested in this case. The first one is that there is a lack of transmission between these two series. Second, we test the hypothesis that the responses of millers to the change in wheat farm price is of the same magnitude regardless of the direction of change (i.e. the price transmission is symmetric). According to economic theory, if a shock arises in a market, its adjustment towards a long-run equilibrium happens via the chain of demand and supply interactions. In a competitive market the rate at which discrepancies from the equilibrium are eliminated is expected to

be the same regardless of whether the initial shock was positive or negative. This would be the case of symmetric price transmission. However, a growing body of literature suggests that the responses of different economic agents to a change in price might be of different magnitude (i.e. asymmetric) between different levels in the supply chain with regards to the direction of change. For example, Peltzman (2000) while analyzing data for 77 consumer and 165 producer goods, concluded that positive asymmetric behavior is the rule rather than the exception. The primary reasons for this are adjustment costs, such as menu costs, or strategic behavior of the intermediaries (Simioni et al. 2013). It would be interesting to see if the same tendency occurs in the Ukrainian wheat supply chain.

Thus, to summarize, five aspects of price transmission are assessed in this paper:

- 1) The magnitude of price transmission, i.e. how much of the price change at one stage of the chain is transmitted to the downwards stage;
- 2) The speed of price adjustment, i.e. the pace at which changes in prices at one level of the chain are transmitted to the other levels;
- 3) The asymmetry of the price adjustment, i.e. the extent to which price increases and decreases are transmitted differently in terms of magnitude and speed to other levels;
- 4) The presence of the structural breaks in the long-run relationship between French export price and Ukrainian wheat farm price;
- 5) The discussion of the factors that could negatively affect magnitude or speed of price transmission in the Ukrainian supply chain with a more detailed focus on export policies.

Analysis of price transmission has become very popular in the agricultural economics literature and has been applied to a number of markets around the globe (for

example, see Simioni et al. 2013, Boetel and Liu 2010, Abdulai 2002). Two of these studies analyze vertical price transmission in the Ukrainian wheat markets. The most recent one is by Gotz et al. (2013) investigates the effects of the wheat export restrictions implemented by the Ukrainian government on the transmission from the world wheat price to the Ukrainian farm level prices during the commodity spike of 2007-08. The study shows that while the series remained cointegrated during the whole period that was analyzed, the magnitude of the price transmission to the growers did change in the 2007-08 crop year. An earlier study by Brummer et al. (2009) studied the price transmission between farm wheat and wheat flour prices in Ukraine from 2000 till 2004. Both of these studies use Markov-switching vector error-correction model in their analysis.

This paper builds on the previous analysis in several ways. First, the data span used in this analysis extends from 2005 till 2012, and looks at the effects of two commodity price spikes (2007-08 and 2010-11) as well as effects of various policy instruments (after a long history of export quotas, in 2011, the Ukrainian government switched to less distorting mechanisms of export tariffs) on the price transmission from the world wheat market into the Ukrainian one. Second, this is the first study that looks into the symmetry of price transmission in the Ukrainian wheat supply chain. Third, the uniqueness of this study in that it tests for the structural breaks in the long-run relationship between world and domestic wheat prices across the entire analyzed period, which was characterized by a variety of events that could have potentially induced such breaks. This, in turn, could provide insights in the causes of structural breaks in the Ukrainian markets, and be used by policy makers in their future decision making process. Finally, in this paper a different set of econometric methods is employed to compare to

those previously used to study this market. This could by itself add to the pool of literature that compares the effectiveness of different cointegration techniques.

2. Background

a. The marketing chain in Ukraine

The major players in the Ukrainian wheat industry are wheat farmers, processors, exporters and retailers. The simplified structure of the Ukrainian wheat supply chain is provided in figure 4.1. Once harvested, most of the wheat is stored at the elevators. From the elevators the wheat can follow the “food path”, i.e. being processed for flour and further into wheat products, the “feed path”, or the “export path”. The leftover of the wheat harvest in a given year is stored in the form of stocks or used for seed in the next year. In 2011/2012, 26 percent of produced wheat was used for exports, 17 percent for domestic feed, and about 52 percent for food and seed (USDA, 2013).

According to the Ukrainian State Statistics Bureau (2013), about seventy eight percent of grain is produced by agricultural enterprises; the rest is produced by the household farms¹¹. Those agricultural enterprises of the size less than 2,000 ha produced eighty five percent of total wheat production in 2011. In the majority of cases farmers, elevators, food/feed processors, and traders are entities independent from each other. However, there are instances when some parts of the supply chain are integrated. Larger farms can have their own storage facilities, or farmers who have their production facilities close to the ports in the South of the country can sell their grain directly to the traders to eliminate the middlemen cost. For example, during the past decade Ukraine (along with Russia and Kazakhstan) has seen an emergence of mega farms, known as

¹¹ Household plots constitute small plots of land (typically less than 0.5 ha), which are used by families for subsidiary farming.

agriholdings. These are the large farms (sometimes larger than 100,000 ha) that are often vertically integrated with processors and/or exporters. According to the Ukrainian Agribusiness Club, in 2011 there were 79 agriholdings in Ukraine with the total land use of 5 million hectares (Gagalyuk, 2011), which is an equivalent of about 25 percent of country's total sown area. Such mega farms account for about nine percent¹² of the total wheat production in the country (Ukrainian State Statistics Bureau, 2013). However, the majority of this wheat is used as fodder for the livestock and poultry produced by the same farms that control about 28% of total animal production in the country (Kobuta et al. 2012).

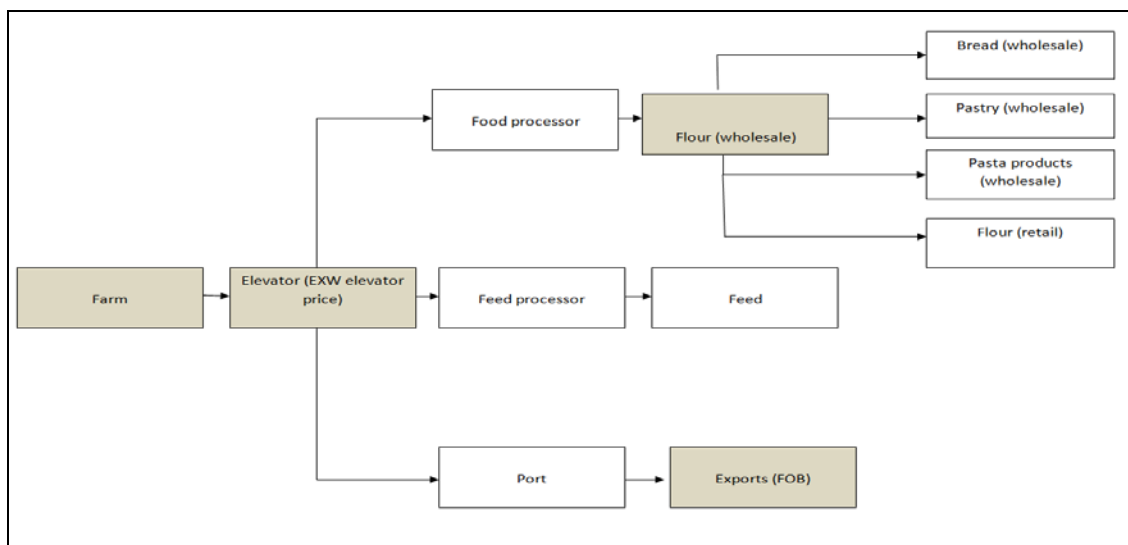


Figure 4.1 Simplified Ukrainian wheat supply chain (without the retail level, except for flour)

Source: Canadian International Development Agency, 2007

The quality of wheat production is dominated by the feed wheat. In accordance with a new adopted standard, feed wheat accounted for 67% in total production in 2009-2010 (Kobuta et al. 2012). Unlike the European wheat quality standards, in the Ukrainian

¹² The estimate is for the farms of the size 3,000 ha or more.

internal market till very recently the quality and class of wheat were determined by gluten content; while in the ports the assessment was based on the protein content of the product. As a result, food wheat (2nd or 3rd class) for the internal market was often classified as feed wheat on world markets.

With regards to the flour production in Ukraine, in 2012 there are about 1500 millers (2012 est.) that are divided into three major categories – industrial (from 100tons/day); agricultural millers (30-100 tons/day), and small miller enterprises (less than 30 tons/day). Seventy percent of the market is controlled by 60-70 large flour producers with the top twenty ones controlling half of the market. Industrial millers produce 70% of the total flour, agricultural millers – 23% and small millers – about 7 percent. Overall, the tendency towards decrease in number and increase in size is persisting in the market (Ukrselko 2013).

The Ukrainian government intervenes in the market by buying grain to support domestic producers; however, the role of state procurement agencies has decreased over time. In 2008-2010 only 5 percent of all grain was sold to them (Kobuta et al. 2012).

Bread prices are controlled by the government. Local administrations can control flour to bread margins and are allowed to set maximum profit margins for bakeries in the country. Such a profit margin usually averages at 5 percent and cannot exceed 10 percent. At the same time, the local administration often provides subsidies to the bakeries to purchase grain and/or flour. Overall, from 2005 until 2012 the real bread price on average increased by 0.7 percent with some fluctuations from year to year. For example, in 2008, the bread price went up by 6 percent compared to 2007, but in 2009 it decreased by 12 percent compared to 2008 (Ukrainian State Statistics Bureau 2013).

b. Export policy environment (2006-2012)

As was mentioned in the introduction, the Ukrainian government is prone to intervene in the wheat export market in a rather ad hoc manner. This section offers a chronology on officially implemented export restrictions in the analyzed period (table 4.1).

Table 4.1 Chronology of grain export restrictions starting from 2006, 1000 MT

Decision date	Period	Wheat	Barley	Corn
10/11/2006	10/17/2006-12/31/2006	400	600	600
12/08/2006	12/14/2006-06/30/2007	3	600	500
02/13/2007	02/15/2007-06/30/2007	3	606	30
02/22/2007	02/26/2007-06/07/2007	3	Quotas cancelled	Quotas cancelled
05/22/2007	05/22/2007	Quotas cancelled	-	-
06/20/2007	07/01/2007-10/31/2007	3	3	3
09/26/2007	01/01/2008-03/31/2008	200	400	600
03/28/2008	04/01/2008 – 04/30/2008	200	400	Automatic licensing
04/23/2008	04/2008-07/01/2008	1,200	900	Automatic licensing
05/21/2008	05/21/2008	Quotas and licenses are cancelled		
10/06/2010	10/20/2010-12/2010	500	200	2,000
12/08/2010	12/2010 – 02/2011	1,000	200	3,000
03/30/2011	04/04/2011 – 07/01/2011	1,000	200	5,000
05/2011	05/2011	Quotas are cancelled		
05/2011	05/2011-01/2012	Tariffs are introduced		
10/2011	10/2011	Tariffs cancelled, except for barley (01/01/2012)		

Source: UkrAgroConsult 2013, Kobuta et al. 2012, FAO and EBRD 2010

In October 2006 in response to rising global grain prices, Ukraine introduced wheat export quota that ranged from 3,000 tons to 1.2 million tons between 2006 and May 2008, when export quotas were abolished in light of an expected extraordinary large harvest. In August 2010, following the Russian ban on wheat exports, Ukraine

implemented a new export quota in the amount of 500,000 tons which was increased to 1 million tons in December 2010. In March 2011, the government announced the extension of the 1 million quotas till July 2011. However, in May 2011, export quotas were substituted with export tariffs that remained in place till October 2011.

3. Methods

Economists have developed a variety of empirical methods to study price transmission (see Fackler and Goodwin 2011), that initially focused on estimating simple correlations between price series and were further developed into much more complex models. In the first step of this analysis, traditional cointegration models, such as the Engle and Granger (1987) procedure (primary one) and the Johansen Maximum Likelihood method (1988) to check for the robustness of the results are used to test for the presence of the long-run cointegration relationship between the analyzed pairs of price series. For the pairs of prices that are found to be cointegrated we use an error correction model, described further in the text to study the short-run dynamics of these prices.

The major limitation of the above methods, however, is that they assume linearity and symmetric adjustment of the price to the long run equilibrium. This might not be an adequate representation of reality and might lead to erroneous conclusions about the price relationships. Therefore, we employ threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) models to account for possible asymmetries in the transmission of the analyzed prices, as well as the Bai and Perron (2003) structural break test to investigate the series for the presence of structural breaks in the long-run relationship.

Consider the Engle and Granger (1987) procedure which consists of two steps. First, the long run relationship between the pairs of analyzed log-prices is estimated¹³:

$$P_t^{FL} = \beta_0 + \beta_1 P_t^{FW} + \varepsilon_t \quad (1),$$

where P_t^{FL} , P_t^{FW} are prices of flour and farm wheat prices respectively. β_0 is an intercept, β_1 stands for the price transmission elasticity, and ε_t is the error term.

In the second step, we use an Augmented Dickey Fuller (ADF) test to check if the null hypothesis that $\gamma_1 = 0$ holds in the following regression:

$$\Delta \bar{\varepsilon}_t = \gamma_1 \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \gamma_{i+1} \Delta \bar{\varepsilon}_{t-i} + \omega_t \quad (2),$$

where ω_t is the white noise term, and $\bar{\varepsilon}_t$ is the residual obtained from the long-run equilibrium equation (1). The number of lags is selected by minimizing the Akaike Information Criterion (AIC) and making sure that errors are not serially correlated. Rejecting the null would mean that analyzed wheat prices are cointegrated, i.e. they move together in the long-run.

While the coefficient β_1 is super-consistent if a pair of prices is cointegrated, it might still be biased for a number of reasons. One of them, described by Stock (1987) and Philips and Loretan (1991), suggests that β_1 is asymptotically biased when two prices are found to be cointegrated because of their correlations in a static equation. Second, if the dynamics of the data-generating process is ignored, it could further lead to the bias in β_1 (Banerjee et al. 1993; Boetel and Liu, 2010).

¹³ From here forward we use an example of the relationship between Ukrainian farm wheat and flour prices. Also for simplicity from this point on Ukrainian wheat farm prices is denoted in the text as “Farm”, while Ukrainian flour price is referred to as “Flour”. French wheat FOB price is referred to as “FrenchFOB”

To correct for the above biases, Stock and Watson (1993) offered a Dynamic OLS technique that adds leads and lags of first differences of the independent variable to the equation (1) that now takes the following form:

$$P_t^{FL} = \theta_0 + \theta_1 P_t^{FW} + \sum_{i=-m}^{+m} \Delta P_{t-i}^{FW} + u_t \quad (3),$$

where P_t^{FL} , P_t^{FW} are prices of flour and farm wheat prices respectively. θ_0 accounts for the transfer costs, θ_1 is a super consistent unbiased coefficient of price transmission elasticity, $\sum_{i=-m}^{+m} \Delta P_{t-i}^{FW}$ is the sum of lags and leads of first-differenced price of wheat, and u_t is the error term. These first differences allow controlling for possible endogeneity and simultaneity problems. Selection of the number of leads and lags in this analysis was based on the AIC. The Newey-West heteroskedasticity and autocorrelation robust standard errors need to be used to assess the significance of the coefficients (Cushman 2000).

a. Asymmetric price transmission

Enders and Granger (1998) argue that the ADF test in equation (2) would be misspecified if price adjustment towards equilibrium is asymmetric. Therefore, they suggested a modification to equation (2) to test for the asymmetric price transmission, which is known as the threshold autoregressive (TAR) model:

$$\Delta \bar{\varepsilon}_t = I_t \gamma_1 \bar{\varepsilon}_{t-1} + (1 - I_t) \gamma_2 \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \gamma_{i+1} \Delta \bar{\varepsilon}_{t-i} + \phi_t, \quad (4),$$

where $\Delta \bar{\varepsilon}_t$ is the first difference of the error term from (1),

$\bar{\varepsilon}_{t-1}$ is lagged error term from (1) lagged for one time period,

γ_1 and γ_2 are the adjustment rates,

I_t is the Heaviside indicator function, such that

$$I_t = \begin{cases} 1 & \text{if } \bar{\varepsilon}_{t-1} \geq 0 \\ 0 & \text{if } \bar{\varepsilon}_{t-1} < 0 \end{cases} \quad (5),$$

and $\sum_{i=1}^p \gamma_{i+1} \Delta \bar{\varepsilon}_{t-i}$ is lagged differenced error term. An AIC is used to determine the optimal lag length.

Alternatively, it is also possible that the process of adjustment to the long run equilibrium may depend on the change in $\bar{\varepsilon}_{t-1}$ instead of the level of $\bar{\varepsilon}_{t-1}$ and, therefore, exhibiting more momentum in one direction than another. For such cases, Enders and Siklos (2001) suggested an alternative momentum threshold autoregressive (M-TAR) model in which Heaviside indicator function is modified in the following way:

$$I_t = \begin{cases} 1 & \text{if } \Delta \bar{\varepsilon}_{t-1} \geq 0 \\ 0 & \text{if } \Delta \bar{\varepsilon}_{t-1} < 0 \end{cases} \quad (6).$$

To summarize, TAR and M-TAR models correspond to “deep” and “steep” asymmetric processes, respectively (Sichel 1993). Deepness corresponds to skewed price series in levels, while steepness represents skewed first-differenced series. However, in either model, $|\gamma_1| < |\gamma_2|$ would suggest that increases in the price of the independent variable tend to persist, while the decreases are transmitted back to the long run equilibrium. The choice of the model is selected based on the AIC or BIC criterion (Enders and Granger 1998). Moreover, several studies (Enders 2001; Enders and Siklos 2001) suggest that the M-TAR model has a greater power in detecting a cointegration relationship than the TAR model.

For both TAR and M-TAR models, the first step is to check (and in case of our paper, to verify) that 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 the price adjustment is symmetric. Rejecting the null, however, would suggest that the series respond differently to whether the departure from the long-run equilibrium is increasing or decreasing.

b. Short-run dynamics

Short-run cointegration tests serve to verify if analyzed prices in different markets respond immediately to the long-run relationship. Error-correction models (ECM) are used to analyze such short-run dynamics. However, their specifications, as in the case of the long-run dynamics analysis, depend on whether price adjustment is symmetric or not.

Symmetric ECM takes the following form:

$$\Delta P_t^{FL} = a_0 + a_1 \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \delta_i \Delta P_{t-i}^{FL} + \sum_{j=1}^n \theta_j \Delta P_{t-j}^{FW} + \vartheta_t \quad (7),$$

where ΔP_t^{FL} and ΔP_t^{FW} are vectors of the first differences of log prices for flour and wheat at the farm level, $\bar{\varepsilon}_{t-1}$ is the lagged residual from (1), ϑ_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).

If, however, series respond with different magnitude towards positive and negative changes (i.e. null hypothesis that $H_0: \gamma_1 = \gamma_2$ is rejected in (4)), ECM model is modified into a threshold ECM (TECM) as follows:

$$\Delta P_t^{FL} = a_0 + \rho_1 I_t \bar{\varepsilon}_{t-1} + \rho_2 (1 - I_t) \bar{\varepsilon}_{t-1} + \sum_{i=1}^p \delta_i \Delta P_{t-i}^{FL} + \sum_{j=1}^n \theta_j \Delta P_{t-j}^{FW} + \mu_t \quad (8),$$

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 threshold¹⁴.

The signs of the short-run adjustment coefficients are expected to be negative. Both equations (7) and (8) can be specified with additional lags ($\sum_{j=1}^n \theta_j \Delta P_{t-j}^{FW}$ and $\sum_{i=1}^p \delta_i \Delta P_{t-i}^{FL}$) to deal with autocorrelation which might be present in the error term.

Additionally, following the procedure used by Ghoshray (2002) we can estimate the number of months (n) it takes flour price series to adjust back to the equilibrium after the change in the farm wheat prices. The formula to use is the following:

$$n = \frac{\log(1-p)}{\log(1-a_1)} \quad (9),$$

where p is a given proportion of the disequilibrium to be corrected, and a_1 is the short-run adjustment speed coefficient from (7) given the symmetric adjustment. If TECM is used, then instead of a_1 in (9), we use ρ_1 and ρ_2 to differentiate between the speed of adjustment to the equilibrium after positive and negative changes.

c. Structural breaks

Since French FOB-Farm and Farm-Flour series are cointegrated, the results of the regressions that analyze the relationships between them are consistent (see equation 3). Thus, β_1 can be considered as the long-run price transmission elasticity. However, such an interpretation of the parameter is grounded in the assumption that the long-run relationship between the series is constant over time, which might necessarily be realistic, especially given the sporadic changes in the Ukrainian grain export policies or two large price commodity spikes in the analyzed period. Thus, we use Bai and Perron (2003)

¹⁴ In this study we focus on testing long run asymmetry only, however, one should note that the model in (10) could further be modified to incorporate short run asymmetries

procedure to test the null hypothesis that there is at least one structural break in the relationship of the French FOB – Farm wheat price pair¹⁵ over the analyzed period.

The pioneer of the structural break tests is considered to be Chow who in 1960 introduced the first test to check for the presence of structural break in the series. The major limitation of his test, however, is that the break date needs to be known *a priori*. This is hard to do in the case of this analysis due to a variety of different events and factors that could have potentially affected the relationship of the analyzed series. By contrast, the Bai and Perron (1998) test (which was further improved in 2003) provides two important advantages over the Chow test and, thus, is suitable for this analysis. First, it allows for searching for more than one break, and, second, there is no requirement for a researcher to know the timing of the break(s) in advance.

The Bai and Perron (1998) procedure is sequential and uses a dynamic programming approach to identify the optimal number of breaks. To do so, at first, a specification of the minimum allowable segment length needs to be estimated. In their 2003 study, Bai and Perron discuss size distortions associated with heteroskedasticity and the autocorrelation consistent covariance matrix estimator when the trimming rate is set at a low level. They suggest a rate at 20 percent for a sample size of 120 observations, yet acknowledge that this rate could be reduced with a smaller sample. It is also not unusual in the literature to set the trimming rate at 0.1 or 0.15 level (Dahl et al.; Jin and Miljkovic). This would imply that each segment tested for the presence of structural break cannot contain less than 10 or 15 percent of total observations. In our case, 10

¹⁵ In this analysis, we test for the presence of structural breaks in the FrenchFOB-Farm relationship only. The reason is that it seems to constitute a more interesting case than a Farm-Flour relationship due to a number of policy interventions in the Ukrainian export market as well as various events in the world grain markets that could have caused such breaks.

percent of the sample would equal to about 42 observations, and 15 percent would account for 63 observations.

Once the trimming parameter is set, the Bai and Perron algorithm evaluates each segment for the presence of structural breaks by carrying out a sequence of F-statistics test for the null hypothesis of no breaks. The optimal number of breaks was selected based on minimizing two criteria - BIC and Residual Sum of Squares. When any extra break is added to the model, its statistical fit improves even if no break was present. Therefore, it is important to focus only on those breaks that improve the model in statistically significant way.

If structural breaks are found in the long-run relationship, the next step would be to re-estimate equation (3) by inserting relevant break dummies in the equation to obtain a consistent estimate of the long-run price transmission elasticity.

$$P_t^{FL} = \partial_0 + \partial_1 P_t^{FW} + \sum_{i=-m}^{+m} \Delta P_{t-i}^{FW} + \mathbf{X}_t + \varphi_t \quad (10),$$

where P_t^{FL} , P_t^{FW} are prices of flour and farm wheat prices respectively. ∂_0 accounts for the transfer costs, ∂_1 is a super consistent unbiased coefficient of price transmission elasticity, $\sum_{i=-m}^{+m} \Delta P_{t-i}^{FW}$ is the sum of lags and leads of the first-differenced price of wheat, \mathbf{X}_t represents regime dummies (here indicated with \mathbf{X}_t for the sake of simplicity), and φ_t is the error term.

4. Application to the Ukrainian grain market

a. Data and stationarity tests

The data used in this analysis consists of weekly nominal French soft wheat FOB prices (Rouen ports), Ukrainian ex-warehouse wheat prices and top grade flour prices. All prices are expressed in logs. The analyzed period spans from January 2005 till December 2012,

accounting for a total of 416 observations. The data was obtained from the Ukrainian analytical agency APK-inform. From visual inspection of the series (Figure 4.2), we can conclude that in general all farm wheat and flour price series tend to move together.

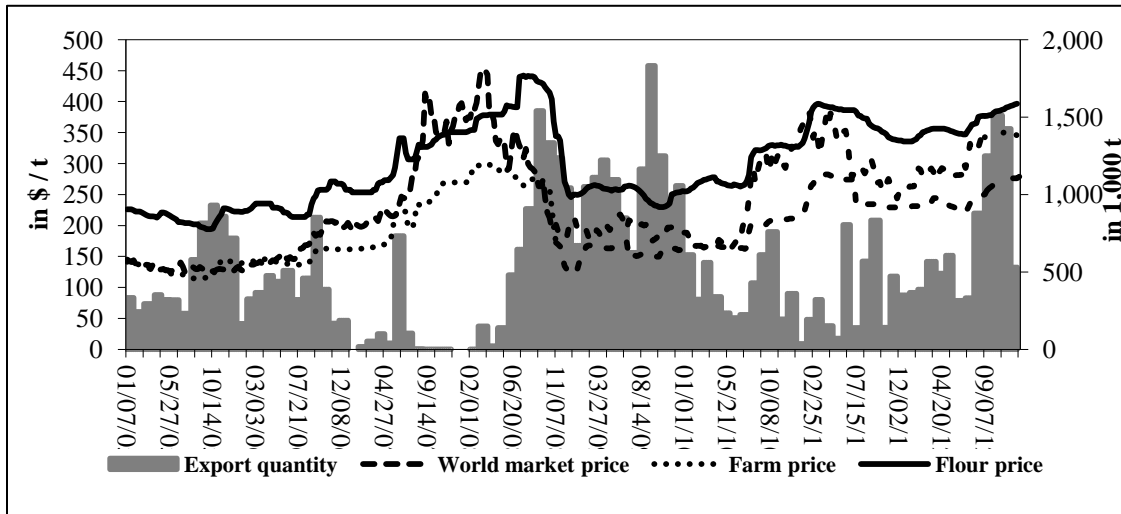


Figure 4.2 Development of the domestic wheat price and wheat exports of Ukraine compared to the world market

Source: GTIS (2013), APK-Inform (2013), HCGA (2013)

However, it can be observed that the domestic prices, even though they share the overall dynamics with the French export prices, do deviate from those at times of the global commodity price spikes of 2007-08 and 2010-11. These periods also coincide with an increased intervention of the Ukrainian government into its wheat exports.

The hypothesis that the price series are non-stationary time series over the whole period were tested by using ADF, PP and KPSS unit root tests. The results suggested that all series are I(1) at conventional significance levels (see table 4.2 and table 4.3).

Table 4.2 Results of the unit root tests in levels^a

	# of lags	AIC	ADF		PP		KPSS	
			w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
FrenchFOB	1	-1532.5	-1.35	-1.66	-1.23	-1.53	2.93**	0.56**
Flour	1	-2270.9	-1.42	-2.02	-0.92	-1.39	3.24**	0.49**
Farm	1	-1986.0	-1.51	-2.11	-0.97	-1.46	2.90**	0.49**

^aAsterisks 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 4.3 Results of the unit root tests in differences^b

	# of lags	AIC	ADF		PP		KPSS	
			w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
FrenchFOB	1	-1526	-13.29**	-13.27**	-17.61**	-17.59**	0.09	0.09
Flour	1	-2264	-9.30**	-9.29**	-11.33**	-11.32**	0.07	0.08
Farm	1	-1978	-9.68**	-9.67**	-11.51**	-11.49**	0.07	0.07

^bAsterisks 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.

b. Cointegration and asymmetric estimations

To test for the presence of the statistically significant long-run relationship between French and farm prices as well as between farm and flour prices, we use the Engle and Granger (1987) two-step procedure. Results provided in table 4.4 show that both pairs of prices are cointegrated¹⁶. The number of lags in equation (1) was selected by using AIC.

¹⁶ We included trend in equation (1) when specifying long-run equilibrium relationship between analyzed pairs of prices. This increased the consistency of the ADF and PP tests' results. KPSS results are significant at 10 percent level in case of FrenchFOB-Farm pair and 10 percent for Farm-Flour pair. However, we conclude that the series are cointegrated since three tests (ADF, PP and Johansen ML) support cointegration presence.

Table 4.4 Engle and Granger cointegration tests for the wheat price series of interest^c

		ADF	PP	KPSS	
	# of lags	AIC	w/ drift and trend	w/ drift and trend	w/ trend
France-Farm	1	-1648.7	-4.85**	-4.04**	0.113*
Farm-Flour	1	-2203.6	-3.41**	-3.45**	0.09**

^cAsterisks 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.

For the consistency check of cointegration, a Johansen ML model was also estimated (see Johansen 1988). Its results confirmed a statistically significant long-run relationship between the series of interest (table 4.5).

Table 4.5 Johansen ML Pairwise cointegration tests for the wheat price series of interest^d

Pairs of series	Ho(H1)	p	Trace	5% CV
France-Farm	r=0(r>0)	2	44.8**	19.99
	r=1(r>1)		2.90	9.13
Farm-Flour	r=0(r>0)	2	30.55**	19.99
	r=1(r>1)		4.64	9.13

^dAsterisks denote levels of significance (* for 10 percent, ** for 5 percent).

According to the results of equation (1), the long-run price transmission elasticities are equal to 0.69 for the French FOB-Farm pair of prices and to 0.74 for the Farm-Flour pair of prices. The price transmission elasticity indicates the percentage change in the price of wheat at one of the stages in the supply chain in response to a one-percent change at another stage.

To assess the robustness of the parameters of the cointegration vector, a DOLS model was specified (see equation (6)). The results are similar to those obtained from equation (1) for each pair of analyzed prices. The long run price transmission elasticity is equal to 0.71 for French FOB-Farm price and to 0.70 for Farm-Wheat pair. Another

benefit of the DOLS model is that the extracted residuals need not to be free from autocorrelation if Newey-West errors are employed.

c. Structural break test results

To check for the presence of structural breaks in the French FOB-Farm long-run price relationship, we employ the Bai and Perron (2003) dynamic programming algorithm. We first used a trimming rate set at 10 percent. This resulted in seven breaks according to BIC and RSS, with many of them being quite minor (according to the SupF plot). Also, since the goal is to further include structural breaks in equation (12), with seven breaks the number of observations per each regime would be quite low, and, thus, might lead to inconsistent results. Therefore, the Bai and Perron (1998) technique was re-run using 15-percent trimming parameter¹⁷.

Figure 4.3 shows BIC and RSS values that correspond to different number of breakpoints shown on the horizontal axis. The optimal number of breaks is selected by minimizing both criteria. As can be seen from this figure, a BIC criterion has its lowest value at three structural breaks, while RSS is minimized at five breaks. To account for both criteria, we select a middle point of four breaks.

¹⁷ We re-ran the model varying the trimming rate between 0.20 and 0.30. Interestingly enough, as the number of breaks decreases with the larger trimming parameter, two breaks are always present – the one in November 2008 and the one in March 2011. However, four breaks allowed us to divide the analyzed period into the regimes with more and with fewer government interventions, and, thus, are more insightful for the purposes of this analysis.

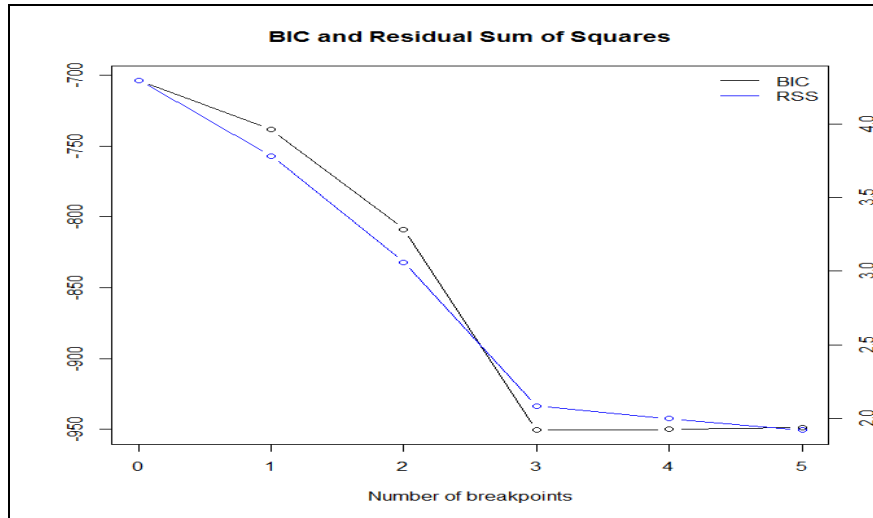


Figure 4.3 BIC and RSS models for m break points

The SupF graph in figure 4.4 supports this choice as F statistics crosses its boundary¹⁸ four times. This can be interpreted as an evidence of four structural breaks at 5 percent significance level.

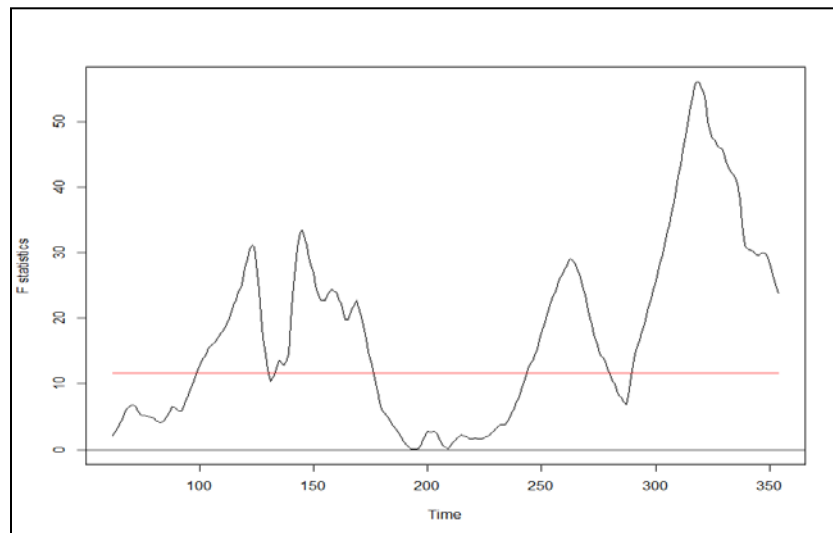


Figure 4.4 SupF test plot for FrenchFOB-Farm pair

Table 4.6 displays the estimates of optimal break dates and their confidence intervals. While it is impossible to be sure what caused these breaks based on the above

¹⁸ Under the null hypothesis, boundary is computed in such a way that the asymptotic probability that the mean of F statistics exceeds its boundary is equal to 5 percent.

results, there are several events that coincided with each break and could possibly explain them.

Table 4.6 Estimated break dates¹⁹

	Break date	Confidence interval	BIC	RSS
Break 1	August 2007	July 2007 – August 2007	2.03	-947.44
Break 2	November 2008	October 2008 – November 2008		
Break 3	January 2010	December 2009 – February 2010		
Break 4	March 2011	March 2011 – April 2011		

Break 1: Figure 4.5 displays the chronology of government interventions in the wheat market over the analyzed period. As can be seen in, in July 2007 Ukrainian government introduced a restrictive quota in the amount of 3,000 tons of wheat due to the severe drought. These months also fall on the first half of the first commodity price bubble of 2007/08.

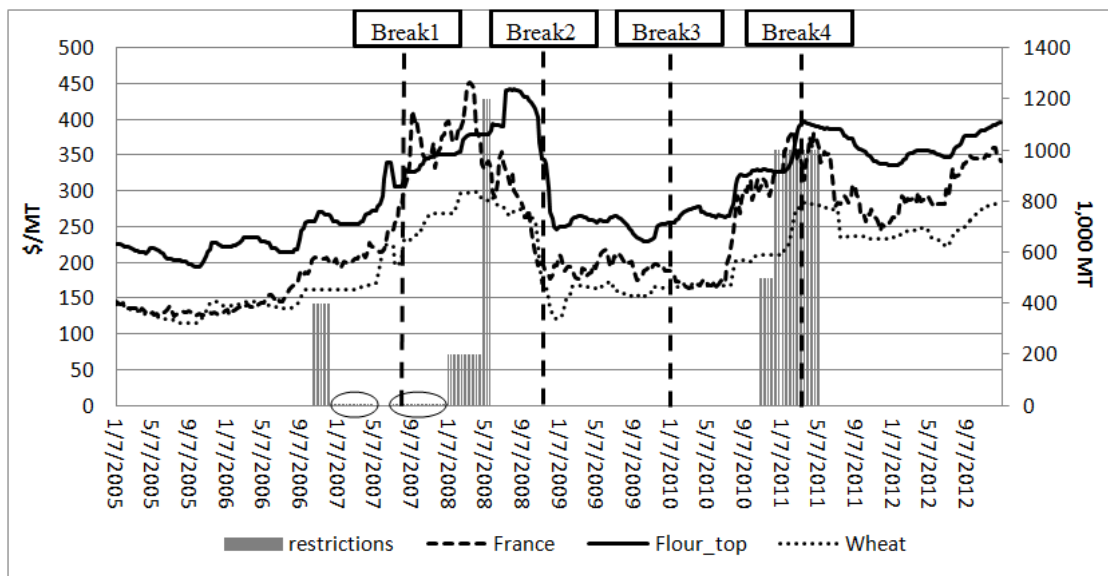


Figure 4.5 Chronology of the export restrictions and structural breaks

¹⁹ If three breaks were to be chosen, they would correspond to August 2007, November 2008 and March 2011.

Break 2: Towards the end of 2008, Egypt, the largest importer of wheat in the world (and one of the largest importers of Ukrainian wheat) had to suspend purchases of Ukrainian product due to complaints about quality issues regarding wheat originating in Ukraine, which resulted in 35% reduction in Ukrainian wheat deliveries to that country as compared to the previous season (Kobuta et al. 2012).

Break 3: This break does not seem to be attributed to a direct policy intervention in the Ukrainian grain market. The exports in early fall of 2009 were among the record ones in history of Ukraine. However, as can be seen in figure 4.2, in January 2010, during the month of the 3rd break, there was almost a 50 percent drop in export amounts from about 1.1 million tons to 600 thousand. This new regime of low export amounts (with maximum export volume equaling to 800 thousand in June and September 2011) remained in the market till the summer of 2012. There are numerous reasons for such a development in Ukrainian wheat export dynamics. Among the major ones are winter kill in early 2010 and persistent heat in the summer of the same year that negatively affected the production of wheat in the country. Such weather conditions were also the major causes for the introduction of the export restrictions in the fall of 2010. Additionally, the president of the Agrarian Chamber of Ukraine named a number of other reasons that caused this decrease in export volumes in 2010. Among them are problems with the VAT repayment, ineffective activities of the Agrarian Fund with grain purchases, and, the most important one – the establishment of the new system of gain classification (APK-Inform 2013).

Break 4: In March 2011, Ukrainian government for the second time in the past three months announced an increase in export quota to 1 mln tons (from previous

500,000 tons). This was also the first time that the government of Ukraine started discussing the imposition of export duties (9% for wheat, 14% for barley and 12% for corn). Later, in May 2011, the quota was, actually, abandoned, and export tariffs were introduced instead till October 2011. Finally, in July 2011, Egypt announces that it considers allowing wheat originating in Ukraine to be included in the next wheat tender (after it has been off the list for 3 years).

Overall, the break dates allow us to divide the entire sample into five periods:

Period 1 (January 2005 till August 2007) is characterized by a tranquil policy situation with no interventions till late 2006, when the first export quota was introduced in the Ukrainian wheat market at the amount of 400,000 tons.

Period 2 (September 2007 – October 2008) is characterized by export quotas restrictions in the market up till May 2008.

Period 3 (November 2008 – November 2010) is another period of a rather tranquil policy situation. Larger quotas (500,000 tons) were introduced in the last month of this time range. This period was, however, also characterized by the suspense of Ukrainian wheat exports to Egypt that could have potentially dampened the price transmission between French and Ukrainian wheat series, since these two countries are major competitors in the Egyptian market.

Period 4 (December 2010 – March 2011) is a short period which coincided with the second commodity bubble, and the beginning of export restrictions imposed by the Ukrainian government.

Period 5 (April 2011 – December 2012) is the period when the government of Ukraine switched to export tariffs, allowing the domestic prices to be more closely linked

to the world prices. These tariffs were abolished in October 2011. Thus, this period could be categorized as the one with the least amount of market distorting interventions.

To summarize, Period 1, Period 3 and Period 5 are the periods with much fewer interventions in the wheat market compared to Periods 2 and 4. Also, the two latter periods coincide with the world commodity bubbles. Therefore, we use the DOLS model with the dummy variables that account for these different periods to test the null hypothesis that the long-run price transmission is not statistically different across different periods. The results provided in Table 4.7, however, suggest that we need to reject the null, and conclude that long-run price transmission elasticity changes significantly in the periods characterized by high governmental intervention and increase in global commodity prices (i.e. period 2 and 4) compared to the periods with lower level of intervention.

Table 4.7 Results of the Dynamic OLS model with regime dummies

Variable	Coefficient ^e
Const	2.17 (0.35)***
Trend	0.0005(0.00)***
p_t^{FR}	0.55 (0.06)***
Regime 1	0.01 (0.07)
Regime 2	-0.34 (0.08)***
Regime 3	-0.33 (0.17)*
Regime 4	-0.14 (0.07)**
Const_r1	-0.08 (0.41)
Const_r2	2.08 (0.48)***
Const_r3	1.6 (0.92)*
Const_r4	0.66 (0.38)*
Δp_{t-1}^{FR}	-0.43(0.09)***
Δp_{t-2}^{FR}	-0.31 (0.09)***
Δp_{t+1}^{FR}	-0.04 (0.09)
Δp_{t+2}^{FR}	-0.02 (0.08)
AIC	-1033.76

^eThe Newey-West heteroskedasticity and autocorrelation robust standard errors are reported in parenthesis. Asterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent).

For example, the elasticity of price transmission is lower by 0.34 in Period 2 compared to Period 5 (the benchmark period used in the model). In Period 4 the elasticity is lower by 0.14 compared to the similar parameter in Period 5. Elasticities in periods 1 and 5 are not statistically different and equal to 0.55. The results for Period 3 are less straightforward. The coefficient for the break 3 dummy is only significant at 10 percent level. However, when a new DOLS model is re-estimated on de-seasoned series, this coefficient becomes statistically insignificant, while other coefficients remain similar in terms of magnitude and significance. Overall, with caution we conclude that there is no significant evidence to suggest that the long-run price transmission elasticity in Period 3 was statistically different from the one in Period 5.

The Ukrainian farmers seem to be the losers from the interventionist policy of the Ukrainian government. If we look at the average difference between French and Ukrainian prices (in absolute terms), it increased twofold in Period 2 and Period 4 compared to Periods 1 and 3 from about 34 USD/ton to 73 USD/ton. Figure 4.6 displays the dynamics of the FrenchFOB-to-Farm price ratios. Rectangles in the figure represent the periods of export quotas implemented by the Ukrainian government. Except for the high ratio in the late 2008 (which was probably attributed to the ban of Ukrainian exports to Egypt as mentioned earlier in the text), it is clearly seen that during the times when quotas were in place the ratio is higher comparing to the periods with no interventions.

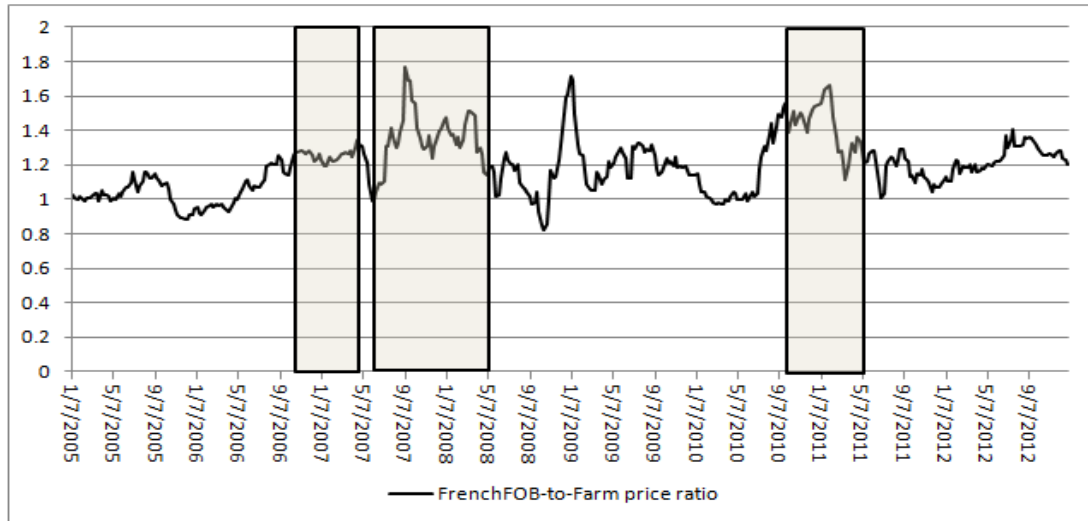


Figure 4.6 Dynamics of French to Ukrainian wheat price ratios

Since choice of the number of breaks was not a straightforward (see figure 4.3), we analyzed the dataset with the inclusion of three breaks only. This resulted in four periods instead of five: Period 1 (January 2005 till August 2007); Period 2 (September 2007 – October 2008); Period 3 (November 2008 – March 2011), and Period 4 (April 2011 – December 2012). Once the dynamic OLS was re-run with the dummies for these periods, the results were consistent with those reported in Table 4.7. Elasticities in more tranquil periods 1 and 5 were not statistically different; elasticity, however, decreased by 20 % in period 2, and by 14% in period 3 that now included export restrictions of the end of 2010, and Egyptian ban of Ukrainian wheat. However, four break division allows us to further disaggregate periods with interventions vs. periods with fewer interventions, and, for this reason, we use four breaks in this analysis.

d. Asymmetry testing

TAR and M-TAR models were estimated next for different lag length to test for the presence of asymmetric price transmission in the analyzed pairs of prices. The models

were run with different lag length, with the optimal number again being selected by the AIC and SBC criterion. The results are reported in Table 4.8.

Table 4.8 TAR and M-TAR model parameter estimates^f

Variable	TAR model		M-TAR model	
	France – Farm	Farm - Flour	France – Farm	Farm - Flour
	Parameter estimate	Parameter estimate	Parameter estimate	Parameter estimate
γ_1	-0.08 (-2.95)**	-0.06 (-2.46)**	-0.06 (-2.01)**	-0.03 (-0.99)
γ_2	-0.06 (-2.37)**	-0.07 (-2.00)**	-0.09 (-3.44)**	-0.10 (-4.33)**
$H_0: \gamma_1 = \gamma_2 = 0(\Phi)$	7.07**	7.48**	7.72**	9.75**
$H_0: \gamma_1 = \gamma_2 (F)$	0.18 [0.67]	0.02 [0.89]	0.81 [0.37]	4.61 [0.03]**
AIC		-1539.70		-2044.16

^fAsterisks 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.

From the results of the TAR test we can conclude that the price transmission between both pairs of prices is symmetric. However, the results of the M-TAR model contradict such a conclusion. The F test in the M-TAR model for the Farm-Flour pair of prices suggests the rejection of the null hypothesis of symmetry. Based on the AIC criterion for both TAR and M-TAR model, it is clear that the latter model fits the data better. Moreover, as was mentioned in the methods part of the paper, the literature suggests that the M-TAR model is better at detecting asymmetry in price transmission among series; however, we use both models for the purpose of consistency check. Based on the results, we conclude that price transmission between farm and flour prices is asymmetric. The coefficients provided in table 4.8 show the level of persistence of positive and negative discrepancies from equilibrium in the short-run. It can be concluded from these results that positive discrepancies in the profit margin of the millers (i.e. wheat price decrease) are not eliminated at all due to the insignificance of the coefficient. The negative discrepancies (i.e. wheat price increase) in the profit margin are eliminated at the rate of 10 percent per week. Therefore, the Ukrainian miller industry responds much

more quickly to the shocks that squeeze its profit margins than to those that stretch them. These results are rather expected, given the structure of the flour market in Ukraine in which more than 50% of all flour production is controlled by twenty largest millers, who can possibly exert some power over the price of their final product.

These results suggest that millers can protect themselves from the increases in wheat prices, by increasing flour prices. However, as the results of the asymmetry tests suggest when the wheat farm prices go down, the millers tend not to pass this decrease onto the bakers. Since the government controls bread prices in the country, abrupt changes in the wheat prices are not necessarily transmitted to the final consumer. This implies, that the major losers in the Ukrainian wheat market are, actually, bakers, who have to deal with controlled bread prices and fluctuating prices in wheat flour or taxpayers, since as was mentioned in the introduction, local government administrations often provide the bakeries with the subsidies to purchase flour and maintain their profit margins as flour price fluctuates.

e. Short-run dynamics

Given the results of the asymmetric price transmission test, the error correction model to test for short-run dynamics for the pairs of prices that are analyzed in the study can be constructed. Since the adjustment between French and Farm prices was found to be symmetric, we proceed with the symmetric error correction model provided in equation (7). For Farm and Flour prices we employ a threshold error correction model offered in equation (8).

Table 4.9 shows the results for both models. It can be observed that the error-correction terms for French FOB prices are negatively significant. This suggests that

Ukrainian wheat farm prices adjust to the long-run equilibrium of French FOB prices. Using equation (9) we estimate that it takes wheat farm price about 37 weeks or 9 months to correct 90 percent of disequilibrium after the change in French export price. This is rather low level of adjustment and could be caused by the inefficiencies present in the market, for example, those caused by the policy interventions or substantially higher (if compared to the US and EU) farm to port transport costs.

Table 4.9 Error-correction model parameter estimates^g

Dependent variable	Independent variable	# of lags	Speed of adjustment, ρ_1	Speed of adjustment, ρ_1	Speed of adjustment, ρ_2
FrenchFOB	Farm	1;1	-0.0240 (-1.33)	-	-
Farm	FrenchFOB	2;2	-0.0638 (-5.70)**	-	-
Flour	Farm	1;1	-	-0.02 (-1.23)	-0.11(-6.08)**

^g t-values are given in parenthesis. ^f Asterisks denote levels of significance (* for 10 percent, ** for 5 percent).

On the contrary, the error correction terms are not significant for the wheat farm price. This implies that French prices do not adjust to correct long-run disequilibria after the change in the farm price, and serves as an indication that French export prices serve is a price leader in its relationship with wheat farm series.

In case of the short-term dynamics for the Farm-Flour pair of prices, the results suggest that flour prices adjust to the long-run equilibrium after a positive change in the wheat price, and 90 percent of such adjustment happens in about 5 months. The short-run adjustment to the decreases in wheat price is found to be statistically insignificant at 5 percent level. In the case of the Farm-Flour price relationship, we assume a unidirectional flow of change from wheat farm price to the flour price. Such an assumption is not surprising given the fact that wheat price comprises about 85 percent of the cost of flour and most likely determines its price.

5. Conclusions and implications

This paper analyzes price transmission along the Ukrainian wheat supply chain from January 2005 until December 2012. The results have important implications for evaluating the performance of the Ukrainian wheat market and efficiency of resource allocation among its different players. More precisely, a number of econometric techniques (such as Engle and Granger (1987) procedure, Johansen ML test (1998) and Error Correction Model) are employed to examine the long- and short-run relationships between French export prices and Ukrainian wheat producer prices as well as between Ukrainian wheat producer and flour prices. Additionally, the assumptions about linearity and symmetric adjustment of the analyzed price to the long run equilibrium are tested with the help of the threshold autoregressive (TAR) and the momentum-threshold autoregressive (M-TAR) models, and the Bai and Perron (2003) structural break test.

Cointegration test results suggest that Ukrainian farm wheat prices are cointegrated with French prices, while flour prices are cointegrated with the farm prices. The estimated long-run price transmission elasticity for the FrenchFOB-Farm pair is equal to 0.70, while for Farm-Flour pair – to 0.71.

The Bai and Perron (2003) procedure suggested that there are four structural breaks in the FrenchFOB-Farm relationship that occurred between 2005 and 2012. Based on these breaks the data was split into five periods. Two of these periods - Period 2 (September 2007 – October 2008) and Period 4 (December 2010 – March 2011) - to a large extent coincided with the world commodity price bubbles and implementation of the export restrictions by the Ukrainian government. Three other periods in between were characterized by a more or less tranquil policy environment with fewer (or less

restrictive) interventions. The results of the Dynamic OLS model applied to the FrenchFOB-Farm pair with the incorporation of dummies for these five periods suggest that during periods 2 and 4, the long-run price transmission elasticity decreased by 0.34 and 0.14, respectively.

As to the price asymmetry tests, they suggest that in case of the Farm-Flour price relationship, increases in wheat prices that result in the reduction of the marketing margin for the millers appear to be passed to the flour consumers faster than reductions in wheat price that lead to increases in millers' profit margin. In case of the FrenchFOB-Farm pair, the price transmission was found to be symmetric.

To analyze short-run dynamics between the prices of interest, error correction models were applied. For the FrenchFOB-Farm pair a symmetric error-correction model was fit. Its results showed that Ukrainian producer prices follow those of the French prices and it takes wheat farm price about 9 months to correct 90 percent of disequilibrium after the change in French export price. If compared to the literature on short-run price adjustment in grain markets (Ghoshray 2007, Dawson et al. 2006), 9 month adjustment to the equilibrium could be considered as a slow rate of adjustment, which is a sign of inefficiencies present in the market. It implies that Ukrainian farmers cannot fully take advantage of a rise in world prices.

To analyze the short-run relationship between farm wheat price and flour price, a threshold error-correction model was employed to account for the found asymmetries. The results suggested that in the short run, after decrease in wheat price, flour prices are irresponsive to the adjustment towards the equilibrium. Increases, however, are passed on within five months, an average adjustment rate. This further suggests that Ukrainian flour

millers might exhibit market power and not pass on decreases in wheat prices to the bakers.

Based on the above results, it could be suggested that during the export restrictions, Ukrainian farmers are among the major losers in the domestic market, since the high world wheat price is not fully transmitted to them. Another category of losers are bakers, who are caught up in-between bread market controlled by the government and millers who are slow at passing on to them the decreases in wheat prices.

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V. EXPORT RESTRICTIONS AND PRICE VOLATILITY IN THE UKRAINIAN WHEAT MARKET: EVIDENCE FROM A DYNAMIC CONDITIONAL CORRELATION GARCH MODEL

1. Introduction

During the recent commodity price booms on world markets (2007/2008 and 2010/2011), export restrictions have widely been used by governments to insulate their domestic markets and prices from price developments on the world market. One such example is Ukraine, a country that has recently emerged from a net importer of wheat into one of its largest exporters. In 2006-08 and later in 2010-11, Ukrainian policy makers used a mix of export quotas and export tariffs to restrict wheat exports. These trade interventions aimed to protect the consumers by preventing the transmission of dramatically increasing and volatile world market prices into its domestic market. By reducing the export quantity, wheat export restrictions increase the supply of wheat on the domestic market which should decrease domestic wheat prices. If this is the case, the consumers can indeed benefit from such policies. However, the producers are the losers, since they cannot take advantage of high world prices. Also, as previous research shows (Goychuk and Meyers 2013), various export restrictions decrease the efficiency of the wheat market, leading to further welfare losses in the economy.

The story with volatility is a bit different, however. Market price volatility has important consequences for the welfare of both consumers and producers, especially if it is caused by adverse and unexpected shocks (Gardner et al. 1977). At the producer level,

high price volatility creates uncertainty and, thus, affects the decision of the farmers to invest. At the consumer level, increases in price volatility translate into larger fluctuations in the purchasing power they hold. By insulating the domestic market from world prices, the government could theoretically protect its internal players from the transmission of high price volatility from the external market. Of course, an important assumption needs to be made here that price volatility is transmitted between different markets in a manner similar to the transmission of price levels by relying on arbitrage processes through trade and information flows. These transmission processes occur until the prices on the domestic and the world market differ at most by the trade costs as suggested by the Law of One Price (Fackler and Goodwin, 2001). Since export controls reduce arbitrage and thus trade flows, prices and their changes in the world market are transmitted less completely to the domestic markets. In this case, both producers and consumers benefit from reduced volatilities, holding everything else constant. And this was one of the justifications used by Ukrainian policymakers when implementing the export restrictions.

However, if export restrictions are implemented on short notice and their design is changed multiple times, as was the case with Ukraine, they could increase marketing risk and might induce additional price variations in the domestic market. As an example, Brümmer et al. (2009) have identified a causal link between market instability and policy interventions. Investigating the wheat market in Ukraine from 2000 till 2004 they find increased residual variance within a Markov-switching vector error correction model in times of ad hoc and frequently uncoordinated nature of domestic policy interventions.

This paper aims at analyzing the dynamics of wheat price volatility in the Ukrainian domestic market. Two specific research questions we are trying to answer are

the following: How did price volatility in the domestic market in Ukraine develop during the export quota system compared to periods of open trade? And how strong was the relationship between the Ukrainian and world²⁰ wheat price volatility during the analyzed period? The results of this analysis would allow us to test whether the goal to implement the export restrictions in order to decrease price volatility in the Ukrainian domestic market were realized.

We address these research questions by investigating the development of price volatility on the Ukrainian wheat market within a multivariate GARCH approach. For comparison, we include the German wheat market, which did not experience export interventions during the food price peaks of 2007/2008 and 2010/2011, as reference case in our analysis.

While the effect of export restrictions on the world market (e.g. Martin and Anderson, 2011; Anderson and Nelgen 2012a; see Sharma 2011 for an overview) and on the domestic market (e.g. Götz et al. 2013a, 2012; Abbott, 2012; Anderson and Nelgen 2012b; Grueninger and von Cramon-Taubadel 2008) has been identified in various studies, their impact on domestic price volatility has not yet been investigated comprehensively. Anderson and Nelgen (2012b) use the standard deviation, the coefficient of variation and the Z-statistic of the domestic price relative to that of the border price as indicators for domestic market instability. The analysis is conducted for 75 countries for all agricultural products for 1955 to 2004. Results suggest that governmental market interventions only slightly increase domestic price stability. Götz et al. (2013) identify an increase in the standard error of domestic prices in Russia and Ukraine during restricted exports within a Markov-switching error correction model.

²⁰ French wheat export price is considered as a measure for the world market price in our models.

They conclude that the export restrictions did not prevent the increase of market instability when compared to Germany and the USA, two countries which did not intervene in their wheat export markets.

Section 2 gives some background information on the export quota system in Ukraine. Section 3 describes our research methods, and the data is presented in section 4. Section 5 gives empirical results, and conclusions are drawn in section 6.

2. Background on wheat trade policy interventions in Ukraine

The government of Ukraine quantitatively limited wheat exports during the two recent commodity price booms both by export quotas which, were implemented within a governmental license system, and export tariffs. Table 5.1 offers a chronology on officially implemented export restrictions in the analyzed period.

In October 2006 in response to rising global grain prices, Ukraine introduced wheat export quotas that ranged from 3,000 tons to 1.2 million tons between 2006 and May 2008, when export quotas were abolished in light of an expected extraordinarily large harvest. In August 2010, following the Russian ban on wheat exports, Ukraine implemented a new export quota in the amount of 500,000 tons, which was increased to 1 million tons in December 2010. In March 2011, the government announced the extension of the 1 million quotas till July 2011. However, in May 2011, export quotas were substituted with export tariffs that remained in place till October 2011.

These trade policy interventions were accompanied by a dramatic increase in political uncertainty since 1) the export quotas were implemented on short notice or in some cases once announced were not implemented at all, 2) export quota sizes were

changed multiple times, and 3) the quota distribution incentivized rent seeking behavior, particularly in 2010/2011.

Table 5.1 Chronology of grain export restrictions starting from 2006, 1000 MT

Decision date	Period	Wheat	Barley	Corn
10/11/2006	10/17/2006-12/31/2006	400	600	600
12/08/2006	12/14/2006-06/30/2007	3	600	500
02/13/2007	02/15/2007-06/30/2007	3	606	30
02/22/2007	02/26/2007-06/07/2007	3	Quotas cancelled	Quotas cancelled
05/22/2007	05/22/2007	Quotas cancelled	-	-
06/20/2007	07/01/2007-10/31/2007	3	3	3
09/26/2007	01/01/2008-03/31/2008	200	400	600
03/28/2008	04/01/2008 – 04/30/2008	200	400	Automatic licensing
04/23/2008	04/2008-07/01/2008	1,200	900	Automatic licensing
05/21/2008	05/21/2008	Quotas and licenses are cancelled		
10/06/2010	10/20/2010-12/2010	500	200	2,000
12/08/2010	12/2010 – 02/2011	1,000	200	3,000
03/30/2011	04/04/2011 – 07/01/2011	1,000	200	5,000
05/2011	05/2011	Quotas are cancelled		
05/2011	05/2011-01/2012	Tariffs are introduced		
10/2011	10/2011	Tariffs cancelled, except for barley (01/01/2012)		

Source: UkrAgroConsult 2013, Kobuta et al. 2012, FAO and EBRD 2010

To give just a few examples: According to the Ukrainian analytical agency UkrAgroConsult (2013), on December 6, 2006 both President Yushenko and Prime Minister Yanukovich informed the Bloomberg News Agency that till the end of December “Ukraine will remove restrictions for grain export introduced in October”. The next week, however, a new Government resolution #1701 was announced that decreased the amount of wheat export quota to a stringent 3,000 tons from the previous 400

thousand tons. In February 2007, the Ukrainian government announced an increase in wheat export quota to 228 thousand tons, however it was never implemented and stayed at 3,000 tons till May 2007. As can be seen from Table 5.1, in June 2007, the Ukrainian government cancelled the export quota for wheat, but already in July it was reintroduced, again at the 3,000 ton level. In August 2010 even though the wheat export quota was not officially implemented, the ships that had been already loaded with wheat and were ready for departure could not leave the harbor, being blocked by the Ukrainian Customs Service (APK Inform, 2013).

Apart from the inconsistency in announcement and actual implementation of the quotas, the decision to grant the license to export was not done in a transparent way. For example, in October 2006 43 traders applied for such a license to export barley. At the end, only five companies (Serna, Suntrade, Reider-Trade, Louis Dreyfus Ukraine and Barge) were granted permission to export more than two thirds of all barley exports (UkrAgroConsult 2013).

3. Methods

In this study we use the Dynamic Conditional Correlation (DCC)-GARCH approach (Engle 2002) to examine and compare the dynamics of volatility of the world wheat prices (represented by the French FOB prices) and domestic wheat prices of Ukraine and Germany. Multivariate GARCH models are common methods used to study volatility in time series. They allow for both analyzing the volatility dynamics of a particular series and investigating volatility correlations and transmissions among several series. More specifically, DCC models are used to approximate a dynamic conditional correlation

matrix that can be used to evaluate the level of interdependency between the series over time.²¹

Consider the following VAR model:

$$y_t = \partial_0 + \sum_{i=1}^p \partial_i y_{t-i} + \varepsilon_t \quad (1),$$

where y_t is a 3x1 vector of French (world), Ukrainian and German wheat price series, ∂_0 is a 3x1 vector of drifts, and ε_t is a 3x1 vector of error terms. ε_t has the following conditional variance-covariance matrix:

$$H \equiv D_t R_t D_t \quad (2),$$

where $D_t = \text{diag}\{\sqrt{h_{jj}}\}$, $j = 1, \dots, J$, is a 3x3 matrix of the standardized disturbance variances from the univariate GARCH models generated for each series. A univariate GARCH (1,1) model can be represented as follows:

$$h_{jj,t} = \gamma_j + \omega_j \varepsilon_{j,t-1}^2 + \delta_j h_{jj,t-1} \quad \text{for all } j = 1, \dots, 3 \quad (3),$$

with $\varepsilon_{j,t-1}^2$ being squared lagged residuals from (1), and $h_{jj,t}$ is a time-varying standard deviation that is further used in defining a GARCH-DCC model. Parameter γ_j is a weighted long run variance. Parameter ω_j indicates the impact of the lagged error term (or, in other words, the role of the previous shocks) on the on series' volatility in a current period. Parameter δ_j represents the effect of price volatility in the previous period on volatility in the current period or decay of volatility over time (Brummer et al. 2009). Both parameters are restricted to be non-negative, and their sum should not exceed unity. One of the weaknesses of the GARCH(p,q) model is that it does not allow for incorporating asymmetry of the errors in the model. To do so, other models need to be

²¹ See Bauwens et al. (2006) for a survey on multivariate GARCH models

considered, such as the EGARCH model of Nelson (1991) or CJR-GARCH model of Golestien et al. (1993).

R_t from (2) is a 3x3 symmetric dynamic correlations matrix of the standardized residuals that is defined in a following form:

$$R_t = (\text{diag}(Q_t))^{-1/2} Q_t (\text{diag}(Q_t))^{-1/2} \quad (4),$$

$$\text{where } Q_t = \{\rho_{ij,t}\} = (1 - \alpha - \beta)\bar{Q} + \beta Q_{t-1} + \alpha(u_{t-1}u'_{t-1}) \quad (5).$$

In the equation (5), $Q_t = \{\rho_{ij,t}\}$ is a time varying covariance matrix of standardized residuals from (1), \bar{Q}_t is the unconditional variance-covariance matrix obtained from estimating a univariate GARCH in equation (3), α and β are non-negative scalars satisfying $\alpha + \beta < 1$, and $u_t = \frac{\varepsilon_t}{\sqrt{h_{jj,t}}}$ is standardized residual from (1). The use of standardized residuals allows for correction of the heteroskedasticity problem when estimating the GARCH DCC model. Parameters α and β can be interpreted as “news” and “decay” parameters capturing the effects of innovations on the conditional correlations over time and their persistence (Hernandez et al. 2012). If $\alpha = \beta = 0$, then the constant conditional correlation model is sufficient, i.e. R_t in (2) can be substituted by R .

The primary focus of the GARCH-DCC model is on obtaining conditional correlations $r_{ij,t}$ in R_t :

$$r_{ij,t} = \rho_{ij,t} / \sqrt{\rho_{ii,t}} \sqrt{\rho_{jj,t}}, \quad -1 < r_{ij,t} < 1, \quad (6).$$

Engle (2002) suggests using a two-step approach to estimate the DCC model by maximizing the following log-likelihood function:

$$L = \left\{ -\frac{1}{2} \sum_{t=1}^T [n \log(2\pi) + \log |D_t|^2 + \varepsilon_t' D_t^{-2} \varepsilon_t] \right\} + \left\{ -\frac{1}{2} \sum_{t=1}^T [\log(R_t) + u_t' R_t^{-1} u_t - u_t' u_t] \right\} \quad (7).$$

The terms between the first brackets are volatility components, between the second ones is the correlation component of the log-likelihood function. Parameters D_t are obtained in the first step and then are used to estimate the correlation component in the second step.

Volatility interdependency could be further tested by performing the Hafner and Herwartz (2006) test and/or Nakatani and Teräsvirta (2009) test. Both tests compute test statistics and associated p-values to test for the causality in conditional variance in GARCH models.

If both test reject the null of no causality, volatility spillover models, such as T-BEKK or EDCC models (see Engle and Kroner 1995; Nakatani 2010), can be specified to study the magnitude of volatility transmission across the series.

4. Data

We conduct our volatility analysis on the wheat market based on 417 weekly observations for the domestic price in Ukraine and Germany, and the world market price from January 2005 until December 2012. We use ex warehouse prices of milling wheat of class III of Ukraine (APK-Inform 2013) and average warehouse delivery price of bread wheat of Germany (AMI 2013) as measures for the domestic wheat price. The FOB price of wheat (French soft wheat, class 1) at the port of Rouen in France, which is the primary harbor through which wheat is exported by the EU (HGCA 2013), serves as the relevant world market price for the Ukraine, and Germany. All prices are real prices and are converted by weekly exchange rates into US \$/t. Figure 5.1 shows the development of these prices in the given period.

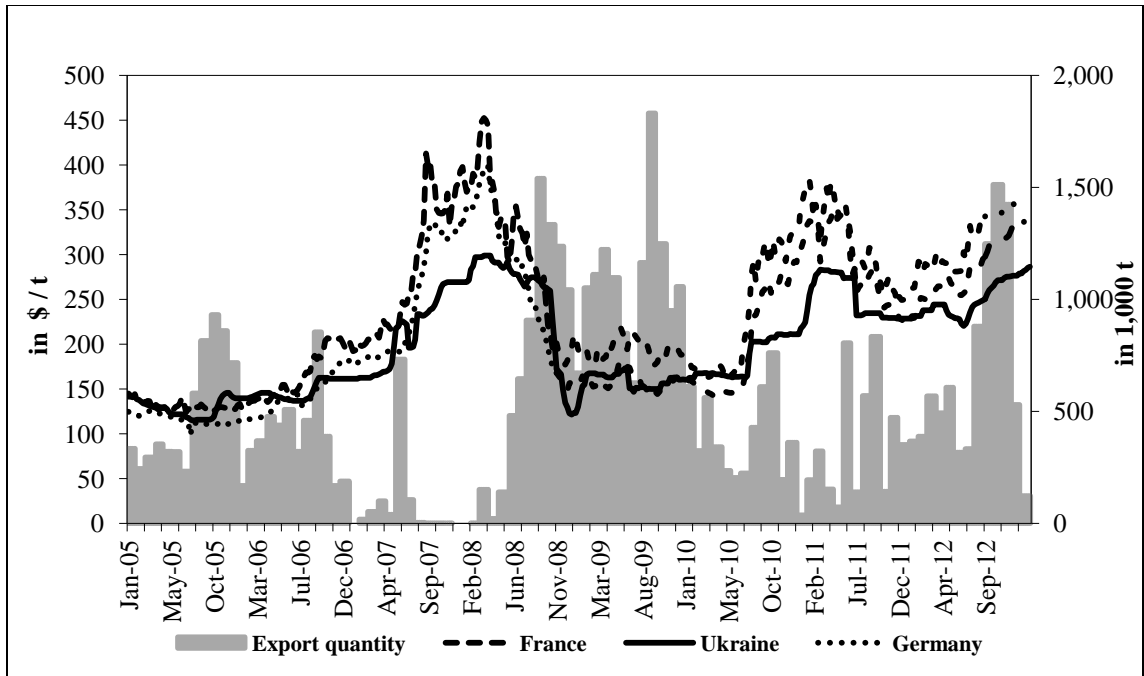


Figure 5.1 Development of the domestic wheat price and wheat exports of Ukraine compared to the world market

Sources: GTIS (2013), APK-Inform (2013), HCGA (2013)

It can be observed that German prices follow closely French prices throughout the entire observed period. However, in the case of Ukraine, the difference between farm and French prices increases during the commodity spikes and impositions of the export quotas by Ukrainian government. Just from visual observation, we could conclude that export restrictions prevented the farmers to benefit from the rising world wheat prices, however, it also saved Ukrainian consumers from further increases in wheat prices. The question this analysis tries to answer is whether the government was able to protect its domestic market from volatility transmission as well.

5. Empirical results

Prior to estimating the VAR and DCC-GARCH model, we check the stationarity of each analyzed series to ensure its appropriateness. All three unit-root 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 (tables 5.2 and 5.3). The results showed that all the differenced series are stationary, leading to the conclusion that the price series of Ukraine, France, and Germany are I(1).

Table 5.2 Results of the unit root tests in levels^a

	# of lags	AIC	ADF		PP		KPSS	
			w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
France	1	-1532.5	-1.35	-1.66	-1.23	-1.53	2.93**	0.56**
Germany	2	-1907.2	-1.32	-1.68	-0.86	-1.20	2.83**	0.50**
Ukraine	1	-1986.0	-1.51	-2.11	-0.97	-1.46	2.90**	0.49**

^aAsterisks 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 5.3 Results of the unit root tests in differences^b

	# of lags	AIC	ADF		PP		KPSS	
			w/ drift	w/ drift and trend	w/ drift	w/ drift and trend	w/ drift	w/ trend
France	1	-1526	-13.29**	-13.27**	-17.61**	-17.59**	0.09	0.09
Germany	1	-1907	-9.43**	-9.42**	-13.07**	-13.06**	0.12	0.12*
Ukraine	1	-1978	-9.68**	-9.67**	-11.51**	-11.49**	0.07	0.07

^bAsterisks 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 5.4 provides some distribution characteristics of the series in first differences of the data in logarithm, i.e. the returns. The mean of the returns is highest for Germany, followed by the world market price and Ukraine. Though, the coefficient of variation indicates that price fluctuations are highest for the French world market price,

followed by the domestic price for Ukraine and Germany. Skewness results indicate that the German price and the French world market price are relatively symmetrically distributed, while prices for Ukraine are less symmetric. Excess kurtosis suggests that the data series are not normally distributed, so that needs to be taken in account when selecting distributions in the following steps.

Table 5.4 Characteristics of the data series in returns^c

	Ukraine	Germany	World
Mean	0.002	0.002	0.002
Standard deviation	0.027	0.027	0.038
Coef. of variation	16.277	11.260	18.625
Skewness	-1.238	-0.141	0.1000
Kurtosis	15.530	4.899	2.715
DF statistic	-6.53***	-5.354***	-6.009***

^cAsterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent)

In the next step we proceed with the Box-Jenkins methodology to determine the order of an ARIMA (p,q) model. Table 5.5 displays some diagnostics statistics for the residuals from the ARIMA (1,0)²² model for each analyzed series. Ljung-Box test for serial correlation of the residuals suggests that residuals are not serially correlated; therefore, the lag structure of the ARIMA models is sufficient. However, the results of the Jarque-Bera normality tests suggest that all models exhibit non normality in residuals.

Table 5.5 Diagnostic test results^d

	Ljung-Box Q(15) Autocorrelation test	Jarque-Bera Normality test	ARCH (12) LM Heteroskedasticity test
World	17.85 (0.27)	103.81 (0.00)***	27.42 (0.007)***
Ukraine	14.01 (0.52)	5731.15 (0.00) ***	21.91 (0.016)**
Germany	21.07 (0.13)	1521.42 (0.000)***	29.61 (0.003)***

^dp-values are in the brackets. Asterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent)

²² The order of ARIMA model was selected based on Schwarz (SBC) criterion. In case of Ukrainian and world wheat prices, SBC was the lowest for the ARIMA (1,0) model. For the German series, the lowest SBC corresponded to the ARIMA (1,4) model, however, due to the lack of significance of the MA(q) coefficients, we found ARIMA (1,0) model to be the most suitable for the German series.

Based on the results of the ARCH-LM test, we concluded that variances of all analyzed series vary over time, and therefore, univariate GARCH (n,m) models needed to be fit for each series. The selection of an appropriate GARCH order for each wheat series under investigation was made in accordance with the minimum AIC and maximum LogLikelihood values up to n,m = 3. The results showed that for all series, GARCH order of (1,1) was the most appropriate one²³.

In the next step we use univariate GARCH results for each series to fit a DCC-GARCH model. Estimated results are provided in Table 5.6.

Table 5.6 Results for the univariate part of the MGARCH model^e

	France	Ukraine	Germany
γ_1	0.00 (0.16)	0.00 (0.00)***	0.00 (0.04)**
ω_1	0.14 (0.02)**	0.74 (0.00)***	0.37 (0.01)**
δ_1	0.83 (0.00)***	0.25 (0.00)***	0.56 (0.00)***
$\omega_1 + \delta_1$	0.97	0.99	0.93
Log	809	1270	1041
Likelihood			

^eAsterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent). p-values are given in parentheses.

For all three analyzed series, GARCH estimates ω and δ are significant at 5 percent level, and their sums are close to one, which is commonly observed with high frequency data, and is suggestive of high volatility persistence. More specifically, a high δ coefficient implies that there is a strong impact of the own-variance on the volatility of the series. A high ω coefficient, on the other hand, means that the series is susceptible to external shocks. As can be seen from the table, the volatility of French and German prices have large own-variance impacts, i.e. high δ and low ω coefficients, which is even

²³ Before determining the order of the GARCH models, we selected the most appropriate distribution based on the AIC, BIC and LogLikelihood criteria. In all the case, Student t-distribution turned out to be the most suitable.

more pronounced for the French export price data. This is, however, not the case with the Ukrainian wheat series. A relatively large ω combined with a low δ suggest that the volatility of the Ukrainian wheat farm prices is rather sensitive to external shocks and that persistence of price volatility is low. These volatility characteristics become also evident in Figure 5.2, which displays conditional variances of the three series. The high susceptibility of the Ukrainian prices to external shocks is reflected in the many pronounced spikes of the conditional variance. However, since persistence of this volatility is low, the conditional volatility quickly returns back to its mean in the aftermath. In contrast, the high persistence of volatility of the world market prices is reflected in the two long-lasting downward paths after the volatility spikes in August 2006 and in July 2010, respectively.

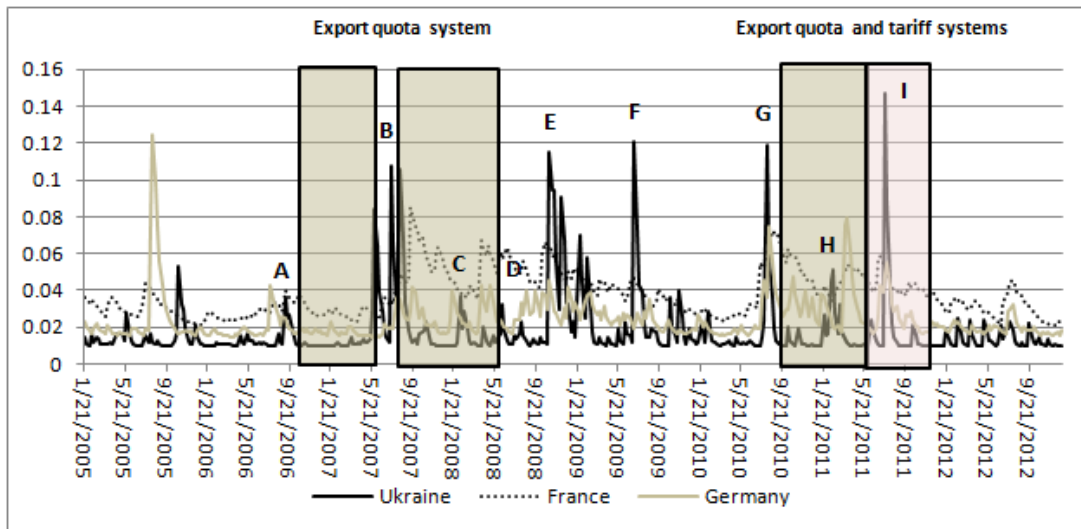


Figure 5.2 Conditional variances of the Ukrainian, German and world wheat market prices

Table 5.7 provides statistics for the distribution of the conditional variances. The mean of the variances is lowest for Ukraine and highest for the world market price. The

same time, standard deviation of the conditional variance is highest for Ukraine, whereas it is lowest for the world market price.

Table 5.7 Characteristics of the conditional variances

	Ukraine	Germany	World
Mean	0.019	0.023	0.038
Standard deviation	0.018	0.011	0.012
Coef. of variation	1.056	2.091	3.167
Skewness	3.860	3.559	0.928
Kurtosis	16.814	19.725	0.424
Minimum	0.011	0.015	0.021
Maximum	0.145	0.121	0.084

Further analysis of the conditional variances provides interesting insights into the development of volatility in the Ukrainian domestic market. Figure 5.2 focuses on the difference among the conditional variances for Ukraine, Germany and France that were retrieved from the univariate GARCH models.

It becomes evident that Ukrainian price volatility is characterized by a number of volatility spikes throughout the analyzed period that exceed the volatility variances of both France and Germany. Figure 5.3 places Ukrainian and world volatility in the context of world price and Ukrainian exports in the analyzed period. It can be seen that world price volatility is rather high during the commodity price booms, particularly during fall 2007 until the beginning of 2009, and again fall 2010 until early summer 2011 during the second phase of high world market prices. However, from visual examination, Ukrainian wheat market did not follow volatility on the world market, rather the majority of volatility spikes happened in between the world price booms.

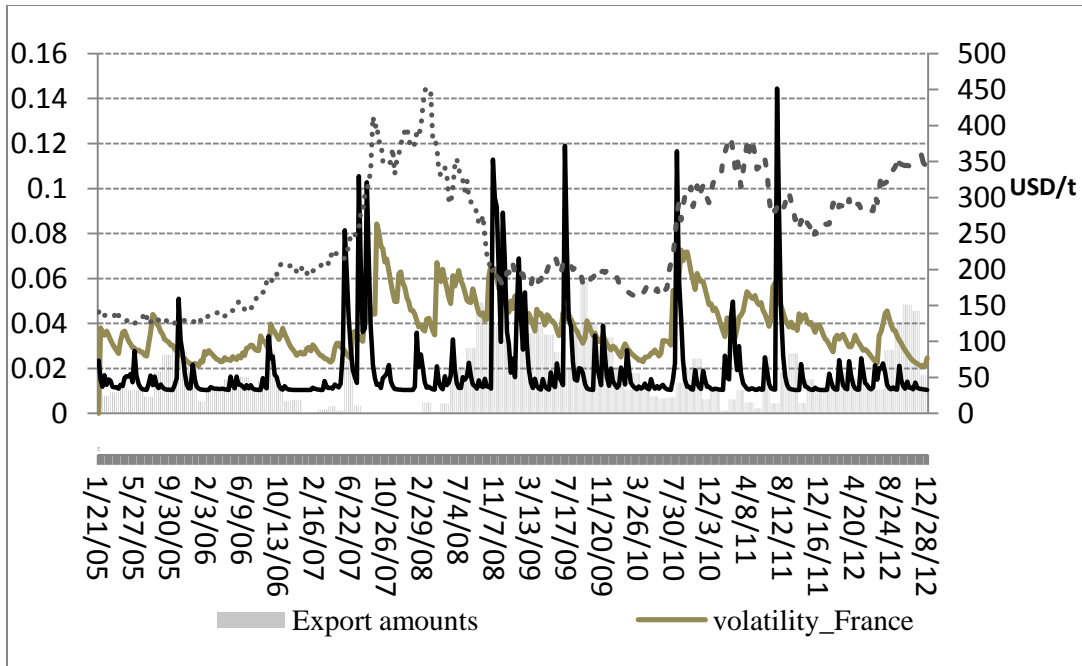


Figure 5.3 Comparison of the Ukrainian and world dynamic conditional variances

Note: dynamic conditional variances (volatilities) are shown on the primary axis; world market price in US\$/t (secondary axis); wheat exports Ukraine in 1000 t

To investigate the possible domestic factors relevant for the wheat price volatility in Ukraine, we identify all political incidences regarding the wheat market in Ukraine for the time periods characterized by excessive price volatility (see Figure 5.2).

Below is the list with political interventions relevant to the wheat market in Ukraine (Götz et al. 2013b):

A: In September 2006 Ukrainian government announces the introduction of export quotas in October 2006, but the size of the quota remains unclear; market experts talk a lot about this in the media.

B: The export quota is lifted on some grains in May (e.g. barley) and for wheat in June 2007; the export quota is reintroduced on July 1, 2007 in light of a severe drought

C: The Ukrainian government announces the increase in the size of the export quota on February 4, 2008 but this is not realized; on Feb 4, the Ukrainian commission on distributing export quotas meets and makes decisions on the export quotas until March 31

D: On April 25, 2008, Ukrainian government prolonged the term of licensing and quotation of grain export till July 1, 2008. However, on May 21, 2008, the government passed the resolution cancelling grain export quotas as of May 23, 2008 (APK-Inform 2013)

E: Towards the end of 2008, the GASC (governmental import company of Egypt) complains about quality issues regarding wheat originating in Ukraine and removes wheat originating in Ukraine from its list (meaning that Ukrainian exporters cannot participate in the wheat tenders)

F: In July 2007 exports of Ukrainian wheat decrease by almost 50 percent compared to May 2007 and by 30 percent compared to June 2007. According to one of the traders in the Ukrainian market, this reduction was caused by “the global financial and economic crisis, as well as deactivation of trading operations” (APK-Inform 2013).

G: Russia introduces a wheat export ban at the beginning of August 2010, which induced discussions in the media whether Ukraine will follow Russia and impose export quotas

H: February and March of 2011 were highlighted by heated debates in the Ukrainian parliament on whether to leave grain export restrictions in place. Towards the end of March, the government introduced the news that in the nearest future export tariffs will be introduced; however, this decision was cancelled the next day. Additionally, in March 2011, Russia announced that it will continue its grain export ban.

I: The Ukrainian government announces the extension of the wheat export quota until the end of June 2011. On June 10, the Ukrainian President signs a law to introduce a wheat export tax on July 1. Towards the end of July, the GASC announces that it considers allowing wheat originating in Ukraine to be included in the next wheat tender (after it has been off the list for 3 years); this was realized for the wheat tender at the end of October

Detailed analysis of the policy environment during the analyzed period makes us conclude that the spikes in volatility of the Ukrainian series coincide with (and are possibly caused by) several types of political events in the domestic market. It should be pointed out that market interventions as such do not necessarily lead to increased volatility. Rather, increased volatility prevails in times of increased market risk which is often caused by political statements and announcements which imply a change in the market conditions. For example, when export quotas are introduced in October 2010, increased volatility cannot be observed in the Ukrainian market. However, a variance spike is observed a few weeks before, when Russia introduced a wheat export ban, which induced the question whether Ukraine would follow Russia in controlling its wheat exports, which was heavily discussed in the media in Ukraine.

The next step of our analysis was to obtain conditional correlations from the DCC-GARCH model (table 5.8). Engle and Sheppard (2001) test results reported in the last line of the table suggest that we reject the null hypothesis that the conditional correlations for each analyzed pair are constant. Therefore, a GARCH DCC model is more appropriate.

Table 5.8 Description of correlation coefficients between two pairs of prices^f

	Ukraine-France	Germany-France
$r_{ij,t}min$	-0.34	0.17
$r_{ij,t}max$	0.39	0.69
$r_{ij,t}mean$	0.05	0.38
χ^2 -test: $R_t = R$	305.5***	442.9***

^fAsterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent). t-values are given in parentheses.

Overall, by looking at the mean values of the dynamic conditional correlations, we can see that the correlation between Ukrainian and French volatility is very low and averages at 0.05. For the case of Germany, it is higher and equals to 0.38. This further suggests higher magnitude of the interdependence between French and German wheat markets vs. Ukrainian and French markets.

Figure 5.4 displays the dynamics of the correlations over time as compared to the conditional variances in French series. Visual analysis of the graph confirms the findings presented in Table 5.9. It can be observed that correlation coefficients for the German-French pair share the co-movement with the French conditional variances, which is not necessarily true for the Ukraine-France pair. This further supports our assumption that the spikes in the Ukrainian volatility are caused by the domestic factors, rather than by the dynamics in the international wheat markets.

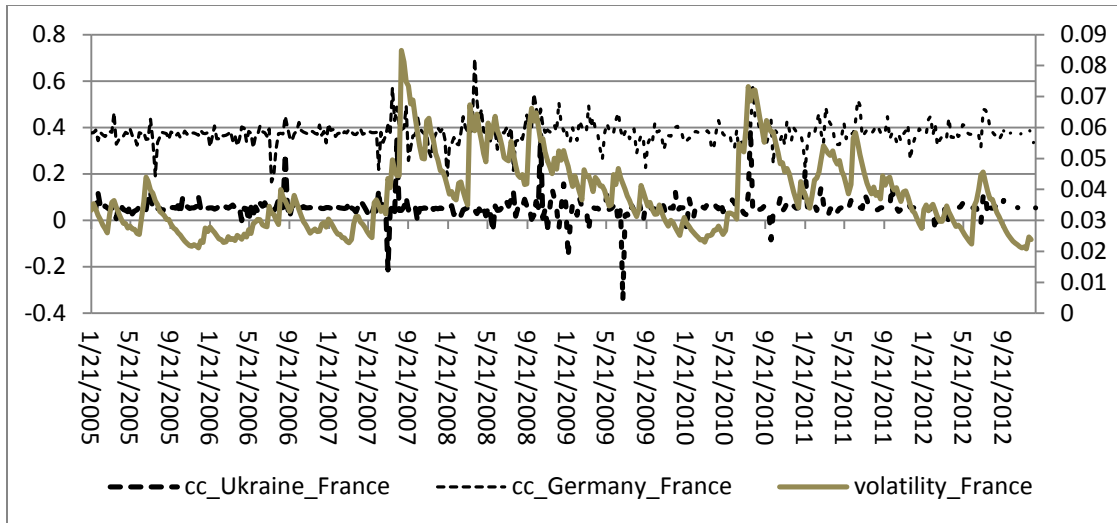


Figure 5.4 Dynamic conditional correlations

Note: “cc” denotes “conditional correlation”

The results of causality in conditional variances tests (table 5.10) show that in case of Ukraine-France pair, Hafner & Herwartz test is significant at 10 percent level, however, Nakatani & Teräsvirta test statistics is not statistically significant. Thus, based on these results we conclude that there is no statistically significant volatility transmission between Ukrainian and French wheat prices. Transmission of volatility between French and German wheat markets is significant at 10 percent level under both tests’ results. Such results confirm our previous findings that volatility in Ukrainian wheat market was not necessarily caused by the volatility in the world wheat market, but was rather determined by the internal factors.

Table 5.9 Results of the causality tests^g

	Nakatani & Teräsvirta test		Hafner & Herwartz test	
	π	p-value	τ	p-value
France-Ukraine	6.04	0.19	8.67	0.07*
France-Germany	9.15	0.06*	9.13	0.06*

^gAsterisks denote levels of significance (* for 10 percent, ** for 5 percent, *** for 1 percent).

6. Conclusions

The empirical results of this study indicate that contrasting to the German and the French world market price, volatility on the wheat market in Ukraine exhibits high susceptibility to external shocks and low impact of own-variance and thus low persistence. This leads to a couple of short periods of time in which excessive volatility prevails, and is reflected in relatively high skewness of the distribution of the conditional variance. Contrasting to the German market, the Ukrainian wheat market did not follow volatility development on the world market 2007/08 and 2010/11, which suggests domestic factors to be of greater importance for observed volatility in this market.

Detailed analysis of the policy environment provides strong evidence for the accordance of phases of high volatility with the occurrence of rumors and the announcement of changes in wheat market trade policy by the Ukrainian government, especially the implementation and extension of the temporary export restrictions.

Further, our empirical results suggest on average lower conditional correlation between volatility in Ukraine and the world market compared to Germany and the world market. We also provide strong statistical evidence for the non-constant, dynamic development of these correlations. The relatively low correlation between the Ukrainian and the world market price volatility compared to Germany can be explained by the high non-tariff trade barriers and high marketing costs in Ukraine. In particular, to export wheat to the world market, a trader has to receive different certificates which cost money and time. Also, due to outdated and insufficient transport and storage facilities, marketing costs are rather high. This implies that the margin between the export and the farmers' price is substantially higher than in Germany, and thus the farmers' price level and

volatility is less closely related to the world market price. However, our analysis does not provide evidence for significantly lower correlation of the Ukrainian with the world market price volatility in times of export restrictions compared to free trade, as suggested by economic theory. Non-significant results of causality in volatility between Ukraine and France further suggest that increased volatility in the domestic wheat market was caused by the internal rather than external factors.

We conclude that the export controls in Ukraine have not significantly reduced price volatility on the domestic wheat market. On the contrary, our findings suggest that the multiple and unpredictable interference of the Ukrainian government on the wheat export market has substantially increased market uncertainty which led to pronounced additional price volatility in the domestic wheat market.

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