EXPORT RESTRICTIONS AND MULTIPLE SPATIAL PRICE EQUILIBRIA WHEN INTERNATIONAL PRICES SPIKE: EXPORT QUOTAS FOR WHEAT IN UKRAINE

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Abstract

Only few models exist which allow for a regime-dependent spatial price equilibrium. This paper focuses on the price insulating effects of export restrictions. The theory of a Walrasian equilibrium and the spatial price equilibrium theory suggest that export restrictions lead to multiple spatial equilibria between the domestic and the world market price. Our analysis is unique in testing for linear versus non-linear cointegration within a smooth transition cointegration model. The application to the wheat export quota in Ukraine shows that the domestic wheat price was stabilized about 30% below the international wheat price during the two recent price booms. We trace back the increased speed of adjustment in the closed trade regime to increased price information flows and heightened information attention when prices are volatile and high. From a global point of view, the domestic wheat price in Ukraine would have increased to the same degree, if no country had engaged in price insulating behaviour 2006-2008 worldwide.

1 Introduction

During the recent price booms on world agricultural markets 2007/2008 and 2010/2011 many countries aimed to insulate their domestic markets from the price developments on the world markets and to stabilize domestic prices by trade policy interventions. Exporting countries have implemented export restrictions reducing or even banning exports, and importing countries have reduced or even completely eliminated import restrictions to dampen the influence of high world market prices on the domestic price level. Most studies on trade policy interventions in the context of commodity price peaks focus on the world market price effects (e.g. Martin and Anderson, 2012; Anderson and Nelgen 2012a; see Sharma (2011) for an overview); there are only few studies which investigate the influences on domestic prices (e.g. Abbott, 2012; Anderson and Nelgen 2012b; Götz et al. 2010; Grueninger and von Cramon-Taubadel, 2008).

This paper develops an empirical framework to identify and measure the effects of export restrictions on the relationship between the domestic and the world market price. How successful were export restrictions in insulating the domestic price from the world market price? Theory suggests that export restrictions reduce not only the speed but also the magnitude of price transmission form the world market to the domestic market. To capture these effects, we develop a flexible spatial price transmission model which allows not only for regime-dependent short-run price transmission but also the parameters of the long-run price equilibrium relationship to change. We apply this model to the wheat market of Ukraine which restricted exports during price peaks 2007/2008 and 2010/2011 by an export quota.

Recent methodological innovations in spatial price transmission analysis have led to increased model flexibility. The linear vector error correction model (VECM) approach was extended, allowing non-linearity in in various respects. For example, the threshold vector error correction (TVECM) and the threshold autoregressive (TAR) models (Balke and Fomby 1997) are motivated by the existence of transaction costs (Goodwin and Piggott 2001). It is assumed that error correction behaviour can be observed only if the deviations from the equilibrium exceed trade costs. As long as the deviations from the equilibrium are smaller than the threshold value, the "neutral band" regime with insignificant or low speed of adjustment prevails. When the deviations from the equilibrium exceed the threshold, an "outside-band" regime characterized by significant error correction behaviour prevails. A smooth transition vector error correction model (STVECM) or smooth transition autoregressive (STAR) model (Teräsvirta 1994) extends the threshold models by allowing for a smooth transition from the "neutral band" to the "outside-band" regime instead of an abrupt regime change in the TVECM and TAR models. This more flexible approach allows inhomogeneity or geographic dispersion of agents which may imply differing transport costs and uncertainty and thus transaction costs (Goodwin et al. 2011; Serra et al. 2006). A Markov-switching vector Error-correction model (MSVECM) is more flexible than the TVECM regarding the threshold variable which induces the regime switches. In the TVECM it is assumed that the size of the deviation from the equilibrium, an observable and deterministic variable, determines the threshold behaviour; in contrast, the regime switches in the MSVECM (Krolzig et al. 2002) are governed by an unobserved and probabilistic variable. Although the above explained price transmission model approaches allow for non-linearity in the short-run price transmission parameters, particularly the speed of adjustment in the TAR and STAR models, they are based on the assumption that a linear and constant long-run price equilibrium exists. However, multiple long-run price equilibria might exist as well with changing long-run price transmission parameters. Only few applications of spatial price transmission models exist which allow for regime-dependent long-run spatial price equilibrium. For example, the magnitude of long-run price transmission changes in the absence of physical trade flows (Stephens et al. 2012), is influenced by governmental market interventions (Myers and Jayne 2011; Götz et al. 2010) and the composition of trade flows (Götz et al. 2008).

Our spatial price transmission analysis is unique in two respects. First, we develop research hypotheses on the effects of the export quota on the relationship between the Ukrainian grower price and the world market price which are based on the theory of spatial price equilibrium (Law of One Price) and the Walrasian equilibrium theory. Second, we introduce a new test on for linear versus non-linear cointegration based on Choi and Saikkonen (2004). We estimate the price transmission parameters in the framework of a smooth transition cointegration model based on Saikkonen and Choi (2004).

2 Wheat Trade Policy Interventions in Ukraine

The government of Ukraine quantitatively limited wheat exports during the two recent commodity price booms by an export quota which was implemented within a governmental license system. Export quotas allow exports up to the amount as specified by the size of the quota. Export quotas varying between 3,000 tons and 1.2 million tons were in force from October 2006 until May 2008 and again from October 2010 until May 2011. Figure 1 shows the development of the Ukrainian wheat grower price (Milling wheat class 3, ex warehouse) and the world wheat market price (French soft wheat, FOB, Rouen) with net wheat exports. Ukraine became a net wheat importer during 2003/2004 when the Ukrainian price increased above the world market price. During times of export restrictions, the domestic wheat price was below the world heat market price and wheat imports were almost not observed.

These trade policy interventions were accompanied by a dramatic increase in political uncertainty since the export quotas were implemented on short notice, the size of the quota was changed multiple times, and their distribution came along with massive corruption, particularly in 2010/2011.

Figure 1 about here

For example, the wheat export quota implemented in 2010 became effective rapidly such that ships already loaded with wheat could not leave the harbour. As a result, several hundred thousand tons of wheat sat in storage temporarily on ships in Ukrainian harbours causing high additional costs to exporters (APK Inform 2010). According to traders' information, this implied that contracts could not be fulfilled, which negatively affected international reputation of traders exporting from Ukraine. Further, the export quota implemented 2006-2008 was first announced in October 2006 to amount 400,000 tons, but it was reduced to 3,000 tons in December 2006. In February 2007 the government gave notice of an increase of the quota to 230,000 tons; however, this increase was not realized. The export quota was abandoned in May 2007 but was reintroduced in June and set at a prohibitive level of 3,000 tons. The notified expansion of the export quota by 200,000 tons in fall 2007 was also not realized. In March/April 2008 the export quota was increased by 1 million tons and finally removed in May 2008 (APK Inform 2010). Also, the majority of the export licenses were distributed to a state owned company in 2010. Foreign grain trading companies did not receive any export licenses unless they paid bribes and thus experienced high economic losses due to foregone exports.

Ultimately, negative incentives for investments are created. Private investments of 550 million € in the grain production sector of Ukraine were delayed in 2010/2011 (EBRD 2011). This has negative effects on global food security since Ukraine has a large grain production potential due to fertile soils and high availability of land, and could further increase its role as a major grain exporter. Together with Russia and Kazakhstan, Ukraine's share in global wheat trade could reach almost 40%, equivalent to more than 50 million tons, in 2019 (Liefert et al. 2010). However, to achieve this private investments are essential since about 1000-2000 US$/ha investments are necessary to fully realize Ukraine's production potential (EBRD 2011).

3 Research Hypotheses

Temporary export restrictions in the context of booming world market prices may influence the relationship between the domestic price and the world market price in various ways. Theory suggests 1) a physical trade effect and 2) a domestic supply effect. We also anticipate 3) the increased availability of market information and strengthened information attention when prices are high and volatile which may also affect this relationship. The following paragraph describes these impacts more in detail.

First, export restrictions dampen the transmission of price changes from the world market to the domestic market via physical arbitrage. According to the Law of One Price, prices on spatially separated markets differ at most by the size of transaction costs of moving goods between these markets. If the markets function well, prices might diverge from this relationship, but the actions of the spatial arbitrageurs bring prices back to their spatial price equilibrium by moving goods from one region to the other (Goodwin and Fackler 2001). If exports are restricted, prices are less transmitted from the world market to the domestic market via the export price, and domestic prices become more determined by domestic market conditions. This implies that the spatial arbitrage condition and the corresponding spatial price equilibrium between the domestic and the world market price may no longer hold and exist. Thus, we expect that the short-run price transmission parameters, particularly the speed of price transmission, decrease. We further anticipate that the magnitude in price transmission, in other words the long-run price transmission, decreases in the longer run as well.

Second, export restrictions change the supply on the domestic market. In particular, the size of the exports decreases, and more is supplied to the domestic market instead, which reduces the domestic market price compared to the price which would prevail if trade was possible. As an example, Figure 2 illustrates the domestic effect of an export quota of the size QA in a partial equilibrium framework. The domestic price p^d equals the world market price p^w minus trade costs TC. The implementation of the export restriction reduces the export quantity q^X to $q^d_{QA} = QA$, the amount as specified by the export quota, and increases domestic supply q^d to q^d_{QA}. This leads to a domestic price p^d_{QA}, which is lower than the price p^d which would prevail if trade was possible, and thus dampens long-run price transmission. The more exports are reduced compared to the open trade regime, and thus the larger the increase of supply on the domestic market, the more does the domestic price decrease. The magnitude of this effect is influenced by several factors. In general, if export restrictions are imposed after the farmer has already decided on his production, the domestic supply elasticity is rather low and thus the domestic price effect is relatively strong. The domestic supply elasticity is also influenced by the degree of warehousing. Particularly larger trading companies with warehousing capacities might not fully supply their products to the domestic market if exports become restricted, but instead keep them in stocks, which increases the domestic supply elasticity, and countervails the above described price decreasing effect of the export restriction. Further, additional policy measures which complement the export restrictions may be of similar effect. For example, governmental purchases on the domestic market, as observed during the wheat export ban in Serbia (compare Djuric et al., 2011), had a price increasing effect on the domestic market. Price increasing effects may also be induced by black trade. Thus, in an extreme case, export restrictions might not prevent domestic prices to increase even beyond the world market price level.

Figure 2 about here

Third, beyond export restrictions, the relationship between the domestic and the world market price might be strengthened by the increased availability of market information in the media and heightened attention to this information in the context of high and volatile world market prices. Besides physical arbitrage, price equilibrating processes could be induced by information and expectations about trade flows (Myers and Jayne 2011)[1]. In the context of this analysis, information on wheat prices prevailing on the world market is spread to domestic grower prices directly by medium and large Ukrainian farmers which have access to up-to-date information on the wheat prices prevailing at the MATIF or the CBOT, and they use this information when negotiating a price with an exporter. Large upward and downward wheat price changes might imply that more market information is supplied through the media and that newspapers report on this issue more often, especially in the case of large price changes with implications for global food security. Also, large price changes might induce heightened attention to market information and actors on the domestic market might be better informed on price developments on the world market. Thus, price changes on the world

[1] Only few papers address the influence of information on spatial price transmission. For example, Stephens et al. (2012) investigate tomato markets in Zimbabwe and find that in the non-trade regime, long-run price transmission and the speed of adjustment are both higher than compared to the trade regime and explain this by information flows between markets (see Literature review below). An another example, Hassouneh et al. (2010) construct a BSE food scare information index based on newspaper articles to assess the influence of the BSE crisis on price transmission among the farm and retail markets for bovine in Spain. They identify strengthened price equilibrating processes in times of the BSE crisis.

market may be transmitted faster to the domestic markets[2] which may imply an increase of the speed of price transmission, given the temporary nature of the export restrictions. Also, price transmission might be observed even in the absence of physical trade flows[3].

Concluding, these theoretical considerations motivate the following hypotheses: First, we expect export restrictions to dampen the magnitude of price transmission in the closed trade regime by a Walrasian price effect and a spatial equilibrium effect (Table 1). Second, if the short-run price transmission parameters, particularly the speed of adjustment, decrease as well, depends on the relative size of the accelerating information effect compared to the decelerating spatial equilibrium effect. Results of previous studies (Götz et al. 2010) suggest the speed of adjustment to increase in the closed trade regime.

Thus, we conjecture that the wheat export quota in Ukraine leads to a regime change not only in the short-run price transmission parameters, but also in the parameters of the long-run price equilibrium relationship between the Ukrainian wheat grower price and the wheat world market price. Therefore, we choose a model framework which captures changes in the short-run as well as the long-run price transmission parameters.

Table 1 about here

4 Literature Review

The effects of wheat export restrictions on the integration of domestic in world markets are also investigated by Götz et al. (2010). They apply a MSVECM to analyse spatial price transmission from the world market to the domestic markets focusing on domestic market effects of wheat export controls in Russia and Ukraine during the 2007/2008 food crisis. They allow for a regime-switch in the long-run as well as short-run price transmission parameters. The model is also applied to Germany and the USA, two countries which did not intervene in their wheat export markets. Three regimes are identified for Russia and Ukraine, with the "crisis" regime prevailing mainly in times of export restrictions. This "crisis" regime

[2] Ghoshray (2011) notices: "…it is possible that world prices are transmitted (to the domestic prices) at a faster rate when prices are unusually higher or depressed than normal; or alternatively, when prices are increasing or decreasing at different rates".

[3] Similarly, Serra et al. (2006) identify price adjustments in the EU pork market even within the "neutral band" of a TAR model, when price changes exceed trade costs, and trace this back to information flows.

is characterized by reduced long-run price transmission which confirms theoretical considerations. Also, an increased speed of adjustment is observed. Comparing results for Russia and the Ukraine with Germany and the USA suggests that the increased speed of adjustment is not caused by the export controls. Instead, it seems that a high speed of adjustment is related with large price changes on wheat markets in connection with intense price information flows.

The main difference of our smooth transition model approach to the MSVECM framework in Götz et al. (2010) is that the regime switches in the long-run equilibrium regression are not abrupt but instead it is assumed that regime switches occur gradually. Furthermore, the MSVECM assumes that the threshold variable is unobservable and probabilistic; however, in our smooth transition model approach the regime switches are assumed to be determined by the level of the world wheat market price. Although the export controls were implemented abruptly on short notice in Russia and Ukraine, some traders might react already in advance before their implementation. When traders expect that world market prices will further increase and rumours come up that export restrictions might be implemented, traders might already change their behaviour. In particular, some traders might increase their export activities temporarily until exports become restricted or forbidden. Other traders might reduce export activities if they expect that exports might be restricted or even banned on short notice in due future to prevent potential losses. Since the point of time at which traders change their behaviour might differ, and traders might behave differently (increase or decrease exports), we assume trader heterogeneity, which justifies the smooth transition cointegration framework. Furthermore, our smooth transition model approach allows explicitly testing for non-linearity in the long-run cointegration regression. In the MSVECM framework in Götz et al. (2010) a LR-linearity test is conducted to test for non-linearity of the model in general.

Goychuk and Meyers (2011) investigate the integration of wheat markets in Russia and Ukraine with world wheat markets based on FOB wheat prices of Russia, Ukraine as wel as USA, Canada and France (2004-2010). Using the Johansen maximum likelihood cointegration test, they find the Russian (but not the Ukrainian) and the French prices to be cointegrated and identify significant short-run and long-run price transmission. They choose a linear model approach whereas we use a non-linear price transmission framework.

Ghoshray (2010) analyses cointegration of international wheat prices within a nonlinear exponential smooth transmission autoregressive (ESTAR)-ECM approach, which allows for nonlinear ESTAR adjustment to the equilibrium (Kapetanios et al. 2006). The choice of this threshold model approach with smooth rather than abrupt change between the “outside-band” and the” neutral band” is motivated by potential variability of transaction costs. Especially, grain transport costs are highly variable and may be influenced by trade volumes. Thus, exporters react different to changes in transport costs. The model is applied to prices observed on world wheat markets in the USA, Argentina, Australia, Canada and the EU. Cointegration is confirmed for all price pairs indicating highly integrated world wheat markets.

The main difference of the model approach in Ghoshray (2010) to our smooth transition model approach is that the long-run equilibrium relationship is assumed to be linear; the behaviour of the speed of adjustment exclusively is assumed to be non-linear, depending on the size of the deviation from the equilibrium.

Gervais (2011) captures nonlinearity in the long-run as well as short-run equilibrium relationship investigating asymmetry in vertical price transmission in the US hog/pork supply chain. The test on smooth transition cointegration by Choi and Saikkonen (2004) confirms non-linearity in the long-run equilibrium relationship and thus asymmetry in the magnitude of price transmission. Parameter estimates indicate downward price stickiness in retail prices. However, results do not indicate asymmetry in the speed of price transmission. Contrasting, previous studies have provided empirical evidence of asymmetry in the speed of price transmission in the US hog/pork supply chain, assuming that the long-run equilibrium relationship is linear. This demonstrates the importance to account for non-linearity in magnitude of price transmission when testing for non-linearity in speed and thus asymmetry in vertical price transmission.

Only few further studies in spatial price transmission exist which allow for non-linearity in the long-run price transmission. Stephens et al. (2012) estimate a regime-specific cointegration model with two regimes distinguishing between price transmission in the presence of trade (“trade regime”) and price transmission in the absence of trade (“no trade regime”) accounting for transaction costs. Their model specification allows the two regimes to differ not only in short-run but also long-run price transmission. They follow Hansen (2003) and use generalized reduced rank regression to test for non-linear cointegration. The model is applied to data on tomato markets in Zimbabwe. Non-linear cointegration is identified, and error correction behaviour is observed in the absence of trade. Even, long-run price transmission and the speed of adjustment are both higher in the no trade regime compared to the trade regime. This provides evidence that spatial price equilibrium might well be established by factors other than physical arbitrage activities of traders, as e.g. information flows. Results further show that if non-linearity in the long-run equilibrium is not accounted for in the model approach, the speed at which prices converge may be underestimated.

Myers and Jayne (2011) investigate spatial maize price transmission between South Africa and Zambia. They propose a multiple-regime threshold model with changing price transmission regimes, allowing for multiple speed of adjustment as well as multiple long-run equilibria. In addition, transfer costs may change over time, thus multiple thresholds may be identified. The regime switches are assumed to depend on the magnitude of trade flows between the regions, temporary governmental market interventions (the government sells imported maize at subsidized prices on the domestic market) and whether transport capacity constraints are binding. When imports are high, cointegration is not confirmed and thus a long-run price equilibrium does not exist. In these periods, the government is the main importer and sells maize at subsidized prices on the domestic market. Thus, the domestic market price decreases, which reduces spatial price transmission, and even breaks off the long-run equilibrium. Results further suggest sensitivity of the price transmission process to the trade volume, especially trade relatively to the size of the market, and the transport capacity. Also, transfer costs are found to vary considerably across time.

Götz and von Cramon-Taubadel (2008) develop a procedure to estimate a regime-dependent VECM which allows not only the short-run but also the long-run price transmission to differ between the regimes. The model approach is applied to apple price data of two German wholesale markets. Non-linear threshold cointegration is confirmed by the Gonzalo and Pitarakis (2006) test, which uses the share of domestic apples in total wholesale trade as the threshold variable. Four price transmission regimes are identified which differ in the equilibrium relationships and the short-run adjustment processes. This model seems to be particularly suitable in settings of irregular seasonal price transmission, typical for fresh fruit and vegetables, in which the use of seasonal dummy variables might not be sufficiently flexible.

Listorti (2009) studies price transmission between the EU and the international soft wheat market (1978-2003) within a non-linear vector error correction model. The influences of the

European Common Agricultural Policy and the implementation of the Uruguay Round Agreement on Agriculture (URAA) are captured by a long-run equilibrium regression allowing for structural breaks in the constant as well as the long-run price transmission parameter. Results suggest the URAA to have strongly improved international price transmission.

5 Data and Empirical Approach

This study uses weekly observations for the world market and Ukrainian domestic wheat prices from March 23, 2001 to September 9, 2011. Ukrainian domestic wheat price is measured as ex warehouse price of milling wheat of class III (obtained from Information Agency APK-Inform). The French soft wheat price (class 1, FOB, Rouen; HCGA 2009) is used as the world market price for Ukraine. World prices and Ukrainian ex warehouse prices are converted based on the daily exchange rates provided by the European Central Bank into US$ per ton. Figure 3 shows the Ukrainian domestic and world wheat price series. Figure 4 presents plots of relationship between these two prices. We also plot the relationships between U.S. and German domestic wheat prices and their corresponding world reference prices as a comparison (there was no active government intervention in grain trade activities by these countries during the two food price crisis periods). Visual inspection leads us to suspect a regime-switching pattern in the relationship between Ukraine and world wheat prices. When the prices are low, the correlation coefficient of Ukraine's wheat price with respect to world reference price is larger than when both prices are high. However, we do not observe such switching behaviour for the U.S. and German situations. This suggests an impact on price linkages resulting from government intervention.

Figure 3 about here

Figure 4 about here

We begin by assessing the time series properties of price series using the standard Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) and the KPSS test of Kwiatkowski, Phillips, Schmidt, and Shin (1992).[i] Table 2 presents the test results. The ADF unit root tests fail to reject the unit root hypothesis for both price series and the KSPP tests reject the stationarity null for the two series. Meanwhile, test results reject the unit root hypothesis and are not able to reject stationarity for the first difference of price series Hence, the price series can be considered as I(1) processes.

Table 2 about here

The next step in the empirical investigation is to estimate the relationship between the Ukrainian and world prices. Introduced by Engle and Granger (1987), the concept of cointegration has become a popular tool in the analysis of nonstationary time series. The premise is that, for two nonstationary I(1) series, if there is a linear combination of them which is stationary, then these two series are said to have a long-run equilibrium and thus are said to be cointegrated. This definition leads to interesting interpretations in the price transmission analysis as the prices can then be interpreted to have a stable long-run relationship and can be represented in a vector error-correction framework.

Empirical implementation involves a two-step procedure for jointly modeling and conducting inferences about the long-run equilibrium together with the short-run adjustment processes towards the equilibrium: 1) estimate the linear equilibrium relationship and test for cointegration; 2) conditional on rejecting the null hypothesis of no-cointegration, test the nonlinearity of residuals, estimate the error correction model (ECM), and investigate how short-run dynamics in the system are influenced by the level, or the sign, of deviations from equilibrium.

Though both economic theories (e.g., market power in supply chain and sticky wage rates in labour markets) and practical economic conditions (e.g., in our case, the state-dependent policy intervention) often imply a nonlinear equilibrium, empirical studies typically only attempt to detect nonlinearity in the adjustment process to the equilibrium while the equilibrium relationship itself has been taken to be represented by a linear regression model.

The development and application of nonlinear cointegrating techniques are still young. Enders and Siklos (2001) propose to test nonlinearity in the residuals of the linear cointegrating vector using threshold behaviour as the alternative hypothesis. The drawback of this approach is that there are no workable approaches to derive a general limiting distribution of this test because the threshold parameters are not identified under the null. Seo (2006) proposes a sup-Wald statistic in the spirit of Davies (1987) to solve the problem, but the procedure is strictly valid only under the assumption that the cointegrating relation is known. Gonzalo and Pitarakis (2006) introduce threshold type nonlinearities within a single equation cointegrating regression model and propose a procedure for testing the null hypothesis of linear cointegration versus cointegration with threshold effects. Krishnakumar and Neto (2009) generalize the estimation and inference procedures of Gonzalo and Pitarakis (2006). However, their threshold cointegrating model requires the threshold/forcing variable to be stationary and ergodic, which may be too restrictive when applying the model to price series, as most of the price data are usually I(1) (Wang and Tomek 2007). For example, in our case, the domestic and world price relationship depends on the world market price, which is a nonstationary series.

Saikkonen and Choi (2004) propose a smooth transition conintegrating (STC) regression model where regressors are I(1) and errors are I(0). The regressors and errors are allowed to be dependent both serially and contemporaneously. Our approach is based on the STC framework of Saikkonen and Choi (2004), Saikkonen and Choi (2004), and Choi and Saikkonen (2010), and follows the procedure suggested by Engle and Granger (1987). The empirical procedures for analysis of the regime-switching price transmission can be described as follows:

1) Test linear versus STC long-run relationship using the method developed by Choi and Saikkonen (2004);

2) Estimate the STC regression model if linearity is rejected in favor of STC (as in our case), using the method proposed by Saikkonen and Choi (2004);ii

3) Test stationarity using the residuals obtained from the estimated STC model;

4) Test linearity versus nonlinearity (e.g., threshold or smooth transition) for error correction procedures, again using the estimated residuals from the estimated STC regression model;

5) Estimate the error correction models, based on the test results from (4), to investigate the dynamic adjustments in the relationship between two prices.

6 Estimation Results

This section aims to explain the theory and to present the empirical results of 1) the test of linear versus STC cointegration based on Choi and Saikkonen (2004), 2) the estimation of the STC cointegration relationship, and 3) the test on threshold cointegration and the estimation of the linear error correction model.

6.1 Linear versus STC Long-run relationship

Consider a smooth transition cointegrating (STC) model

$$(7)\quad y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + z_t,\ \ t = 1, 2, ..., T$$

where y_t denotes the (logarithmic) Ukrainian wheat price and x_t represents the (logarithmic) world reference price; z_t is a zero-mean stationary error term, α_1 and α_2 are constant terms; β_1 and β_2 are parameters that measure the price transmission elasticity, and $g(x_t - c; \gamma)$ is a smooth transition function of the process x_t, with smoothness parameter γ and threshold value c. The non-linear nature of the model is determined by the transition function. Like other smooth transition autoregressive (STAR) models, the STC can be thought of as a regime-switching model that allows for two regimes, associated with extreme values of the transition function, $g(x_t - c; \gamma) = 1$ and $g(x_t - c; \gamma) = 0$, and where the transition from one regime to the other is smooth. The regime that occurs at time t is determined by the observable variable x_t and the associated value $g(x_t - c; \gamma)$. Different choices for the transition function give rise to different types of regime-switching behaviors. In our study, we use a first-order logistic function as the transition

$$(8)\quad g(x_t - c; \gamma) = \left[1 + \exp(-\gamma(x_t - c))\right]^{-1}.$$

The parameter c can be interpreted as the threshold between the two regimes, in the sense that the logistic function changes monotonically from 0 to 1 as x_t increases. When x_t is small (relative to the threshold c), g approaches 0 and the behavior of y_t is given by $\alpha_1 + \beta_1 x_t + z_t$. Similarly, as x_t becomes large, g goes to 1 and the behavior of y_t is then given by $(\alpha_1 + \alpha_2) + (\beta_1 + \beta_2) x_t + z_t$. The parameter γ determines the smoothness of the change in the value of the logistic function and, thus, the smoothness of the transition from one regime to the other. As $\gamma \to 0$, the STC model becomes an AR(p) model. When $\gamma \to \infty$, the regime-switching from 0 to 1 becomes instantaneous at $x_t = c$. Hence, the STC model in (7) includes a two-regime threshold autoregressive (TAR) model as a special case. In the Logistic STC model, the two regimes are distinguished by small and large values of the transition variable

x_t (relative to *c*). This type of regime-switching is appropriate in our case, as the relationship pertains to the active or inactive state of policy intervention, which itself is triggered by the level of world market prices. For detailed discussions on the choice of transition functions, the reader is referred to vn van Dijk, Teräsvirta, and Franses (2002) and Teräsvirta, Tjøstheim, and Granger (2010).

Testing linearity against the STC specification constitutes a first step towards building the STC models. The null hypothesis of linearity can be expressed as equality of the autoregressive parameters in the two regimes of the STC model in (7). That is, $H_0 : \alpha_2 = \beta_2 = 0$, whereas under the alternative hypothesis of H_1: $\alpha_2 \neq 0$ or $\beta_2 \neq 0$. The testing problem is complicated by the presence of unidentified nuisance parameters under the null hypothesis. Informally, the STC model constrains parameters which are not restricted by the null hypothesis, but about which nothing can be learned from the data when the null hypothesis holds. The null does not restrict the parameters in the transition function γ and *c*, but when H_0 holds, the likelihood is unaffected by the values of γ and *c*. Another attractive alternative might be testing the null hypothesis $H'_0 : \gamma = 0$ directly from Equation (8). However, under H'_0, the magnitudes of α_2 and β_2 are completely irrelevant. In other words, the values of α_2 and β_2 are unidentified under the null hypothesis when the model is linear. In this case, it is impossible to perform an LM linearity test. Luukkonen et al. (1988) and Granger and Teräsvirta (1993) develop tests that circumvent the problem associated with the presence of nuisance parameters by replacing the transition function with a Taylor series approximation. However, since we are working with cointegrating regressions, and thus with *I*(1) data, this brings about notable new challenges to the testing problem.

Choi and Saikkonen (2004) develop a nonlinearity test that extends the approaches developed by Luukkonen et al. (1988) and Granger and Teräsvirta (1993), and that can be applied in the context of STC. In particular, their test relaxes the exogeneity requirement for the regressors and follows the common practice in cointegrating regressions and permits both serial and contemporaneous correlations between the regressors and the error term of the model. In order to allow for this feature, the test uses the leads-and-lags approach proposed by Saikkonen (1991) and Stock and Watson (1993) for linear cointegrating regressions.

Following Luukkonen, Saikkonen, and Teräsvirta (1988), Choi and Saikkonen (2004) propose a set of tests based on the first- and third-order Taylor series approximation of the transition function g. The authors argue that a third-order Taylor expansion is superior to a first-order version, since it has more power when β_2 in (7) is small. We thus adopt the third-order Taylor approximation and rewrite the transition function as

$$(9)\quad g(x_t - c;\gamma) \approx b\gamma(x_t - c) + d\left[\gamma(x_t - c)\right]^2 + h\left[\gamma(x_t - c)\right]^3 .$$

The testing procedure involves estimating the corresponding auxiliary regression using OLS[iii]

$$\begin{aligned} y_t &= \alpha_1 + \alpha_2\left\{b\gamma(x_t - c) + d\left[\gamma(x_t - c)\right]^2 + h\left[\gamma(x_t - c)\right]^3\right\} \\ (10)\qquad &+ \beta_1 x_t + \beta_2 x_t b\gamma(x_t - c) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} \\ &= \omega + \phi_1 x_t + \phi_2 x_t^{\,2} + \phi_3 x_t^{\,3} + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \qquad t = K+1,\dots,T-K \end{aligned}$$

The null hypothesis of linearity is $\phi_2 = \phi_3 = 0$ and the LM statistic is $\tau = \hat{\Phi}'[\hat{\sigma}_e^2 (M^{-1})_{xx}]^{-1}\hat{\Phi}$, where $\hat{\Phi} = [\hat{\phi}_2\ \hat{\phi}_3]'$ are the OLS estimates of $[\phi_2\ \phi_3]$, $\hat{\sigma}_e^2$ is the variance estimator based on the residuals of the corresponding OLS estimation constrained by $\phi_2 = \phi_3 = 0$, M is the sample moment matrix for the auxiliary regression, and thus $(M^{-1})_{xx}$ is the element of the inverse of the sample moment matrix associated with $[x_t^2\ x_t^3]'$. Under the null hypothesis, $\tau \xrightarrow{d} \chi^2(p+1)$, where *p* (= 1 in our case) is the dimension of the model. Test results are presented in Table 3.

Table 3 about here

Under all levels of lags and leads (*K*), the test rejects the null of linearity in favor of the STC framework. We thus use the STC for modeling the long-run relationship for Ukraine and world wheat prices. As a comparison, we also test the linearity of the U.S. and German wheat prices with their corresponding world price relationships. Neither of the tests is able to reject the linearity assumption, which suggests STC is not appropriate for the U.S. and German wheat markets. This is consistent with our priors, since these two countries have not implemented trade restrictions during the food crisis. In our next step, we estimate the STC relationship for the Ukrainian case. Of course, as always, before we can draw any formal conclusion about the long-run equilibrium, we will need to test the stationarity of the residuals to decide if indeed these prices are cointegrated.

6.2 Estimation of the STC Long-run Relationship

Given that the null hypothesis of linearity has been rejected, our next step is to estimate the STC model. Previous studies (for example, van Dijk, Teräsvirta, and Franses 2002 and Enders 2010) usually suggest using a nonlinear least square (NLLS) technique to obtain the estimates of the parameters in (7).[iv] The estimate of the parameter vector $\theta = [\gamma\, c\, \alpha_1\, \alpha_2\, \beta_1 \beta_2]$ will satisfy

$$(11) \quad \hat{\theta} = \underset{\theta}{\arg\min}\, Q_T(\theta) = \sum_{t=1}^{T} [\tilde{\ } \text{[illegible]} \vartheta)]^2$$

where [illegible] is sample observations and $y_t(x_t;\theta)$) is the so-called skeleton of the model given in (7). As before, we are working with the STC model where regressors are $I(1)$ and errors are $I(0)$, and the regressors and errors may be dependent both serially and contemporaneously. Saikkonen and Choi (2004) point out that, although the nonlinear least squares estimator from (11) is consistent, the asymptotic distribution involves a bias if regressors and error are dependence, which makes the above NLLS estimator inefficient and unsuitable for use in hypothesis testing. They thus propose a Gauss–Newton (G-N) type estimator that utilizes the NLLS estimator obtained from (11) as an initial estimator and expands the model by including leads and lags as extra regressors. Using leads and lags enables the G-N estimator to eliminate the bias and have a mixture of normals distribution in the limit, thereby making it more efficient than the NLLS estimator and thereby suitable for use in hypothesis testing. That said, the estimation procedure is comprised of two steps: to compute the NLLS estimator $\hat{\theta} = [\hat{\gamma}\, \hat{c}\, \hat{\alpha}_1\, \hat{\alpha}_2\, \hat{\beta}_1\, \hat{\beta}_2]$ for equation (11) and then to use $\hat{\theta}$ as the initial value and estimate the following augmented STC model

$$(12) \quad y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta \ , \ t = K+1, \ldots, T-K$$

The Saikkonen and Choi (2004) approach has provided us with valuable suggestion for obtaining a consistent and unbiased estimates for the STC models. Actually, all methods for nonlinear optimization are iterative: from a starting point θ_0 the method produces a series of vectors $\theta_1, \theta_2, \ldots$ which (hopefully) should converge to θ^*, a global minimum for the given function. If the given function has several (local) minima, the result will depend on the starting point θ_0. Thus, the starting point for estimation is important in the empirical procedure. The Saikkonen and Choi (2004) approach provides a suitable starting point for the second stage G-N estimation. Given the estimate from the first NLLS stage is the true θ^* for the first NLLS estimation, the second G-N approach supplies the better estimates. We adopt their iterative estimation procedure and utilize a damped G-N method, known as the Levenberg-Marquardt (L-M) method. Given the initial values of the parameters are close to the final optimal values, the L-M method has proved to be more efficient and can almost always guarantee quadratic final convergence.

Also, as just discussed, the estimate results could be sensitive to the initial values of γ and c. van Dijk, Teräsvirta, and Franses (2002) thus suggest normalizing the transition function by dividing γ by the sample standard deviation of the transition variable x_t to make γ approximately scale free. We thus replace the transition function (8) with a normalized version

$$(13)\quad g(x_t - c;\gamma) = \left[1 + \exp(-\frac{\gamma}{\hat{\sigma}_x^{\,2}}(x_t - c))\right]^{-1}$$

Table 4 presents the (iterated) L-M estimates of the cointegration models for the linkages between Ukrainian and world wheat markets. Before discussing the results, we need to test the stationarity of the residuals first. We thus conduct a stationarity test utilizing the residuals obtained from the above STC regression. The results are presented in Table 5.

Table 4 about here

Table 5 about here

The KPSS test does not reject the null of stationarity at a 5% level. We also conduct the ADF unit root tests for the residuals. However, since the residual variance is made as small as possible, the procedure is prejudiced toward finding a stationary error process. Hence, the test statistic used to test the unit root must reflect this fact and an ordinary ADF-table is inappropriate. We thus use the critical values provided by Enders (2010, which are interpolated using the response surface in MacKinnon 1991. The results reject the null of unit root. We therefore conclude the Ukraine and world market wheat prices are cointegrated via a smooth transition mechanism.[v]

The STC results from Table 4 are consistent with the institutional background and with our conceptual framework. When comparing the results from STC models with and without lags and leads, we find no significant difference. This may indicate that regressorerror dependence is not an issue in our sample set. Equation (15) is based on the STC with no lags and leads. It reveals the STC long-run equilibrium relationship for the two prices.

$$(14)\quad \hat{y}_t = \begin{cases} -0.86 + 1.14x_t, & \text{if g=0} \\ 1.27 + 0.70x_t, & \text{if g=1} \end{cases}, \quad \text{and } g(x_t - c; \gamma) = 1/\{1 + \exp[-3.87(x_t - 5.21)/0.16)]\}$$

.

The results confirms a regime-switching behavior in the long-run relationship between Ukrainian domestic and world market prices, based on the level of world prices. The estimated threshold value for the transition variable is 5.2 in logarithms, or \$185. When the world price is below the threshold of \$185/ton, the transmission elasticity of domestic price with respect to the world price is about 1.1. The two markets are closely integrated. This provides evidence that when the world price is not "too high", no active export control has been triggered, and thus that price changes or shocks in the world market can be fully transmitted to the Ukrainian market. At the same time, when the world market is "too high" (from the perspective of the Ukrainian government), and exceeds the threshold level of \$185, the relationship between the two markets gradually switch to another regime and the transmission elasticity decreases to 0.70. This reflects the effects of trade interventions on price transmission. The two food crisis periods, with strict export controls, belonging to this regime. The fitted price relationship is also presented in Figure 5. Finally, it is quite interesting to see what happens when the wheat price is between two regimes.

In that case an increase of one unit in the world market price will only partially be passed along to the domestic market while a similar decrease in the world price will fully be transmitted to the domestic market. The domestic growers under such a situation are thus worse off from price increases as compared to the potential benefit they might gain from the same price increase in the world market, all else being equal.

Figure 5 about here

6.3 Estimation of the Error-correction Model

The transaction cost version of the LOP provides justification for using the momentum threshold autoregressive (M-TAR) or Exponential STAR types of regime-switching models which allow the adjustment behavior to be asymmetric inside and outside the transaction cost band. A standard two-parameter and three-regime M-TAR model when applied to the deviations from equilibrium, can be expressed as

$$(15)\quad \Delta z_t = \begin{cases} \phi_1 z_{t-1} + \varepsilon_1, & \text{if } z_{t-1} < \theta_1 \\ \phi_2 z_{t-1} + \varepsilon_2, & \text{if } \theta_1 < z_{t-1} < \theta_2 \\ \phi_3 z_{t-1} + \varepsilon_3, & \text{if } z_{t-1} > \theta_2 \end{cases}$$

where z_{t-1} is the previous deviation from long-run equilibrium. An equivalent vector error correction representation of (16) can be written as

$$(16)\quad \Delta y_t = \begin{cases} \sum_{i=1} \alpha_{1i} \Delta y_{t-i} + \sum_{j=1} \beta_{1j} \Delta x_{t-j} + \varphi_1 z_{t-1} + e_1, & \text{if } z_{t-1} < \theta_1 \\ \sum_{i=1} \alpha_{2i} \Delta y_{t-i} + \sum_{j=1} \beta_{2j} \Delta x_{t-j} + \varphi_2 z_{t-1} + e_2, & \text{if } \theta_1 < z_{t-1} < \theta_2 \\ \sum_{i=1} \alpha_{3i} \Delta y_{t-i} + \sum_{j=1} \beta_{3j} \Delta x_{t-j} + \varphi_3 z_{t-1} + e_3, \quad , & \text{if } z_{t-1} > \theta_2 \end{cases}$$

In (16) and (17), the interval $[\theta_1, \theta_2]$ defines an asymmetric transaction cost band within which arbitrage is not profitable. The ϕ_i can be interpreted as the speed-of-adjustment parameter. In this specification, deviations from the long-run cointegrating relation trigger error correcting movements in prices when the deviations fall outside of the band. If $z_{t-1} < \theta_1$ or $z_{t-1} > \theta_2$, then error correction follows a stationary AR(1) process and trade or arbitrage between markets is profitable. However, we are investigating a situation which is one-sided because of the nature of policy interventions. There is no transaction cost band, only one-sided transaction costs for trade from the domestic market to world market, it is thus more appropriate to utilize a two-regime threshold model to investigate the error correction process.[vi]

We begin by conducting a linearity test for the residuals which is based on Hansen's (1999) self-exciting threshold autoregressive (SETAR) approach. SETAR models with one regime (which shrinks to a linear AR model) and two regimes are

(17) $z_t = \alpha_1 z_{t-1} + e_t$ and

(18) $z_t = \alpha_1 z_{t-1} I_1(\gamma) + \alpha_2 z_{t-1} I_2(\gamma) + e_t$, respectively

where z_t here is the predicted residuals from STC regression, $I(\gamma)$ is an indicator that $I_i(\gamma) = 1$ when ith regime occurs and γ is the threshold. The estimates of α_1 and α_2 are obtained from OLS along with the sum of squared residuals, denoted as SSR_2. The threshold has been chosen when the estimation of (19) gives the minimum sum of squared residuals ($SSR_2^{\min}$), alternatively, $\hat{\gamma} = \operatorname{argmin} SSR_2(\hat{\gamma})$. The search over all possible values of the threshold is restricted to the values of z_{t-1} that lie between the 15th and 85th percentiles. Let SSR_1 denote the sum of squared residuals from (18) and $SSR_2^{\min}$ denote the minimum sum of squared residuals from (19), which is the chosen threshold model, and the *F*-statistic can be constructed as

(19) $F_{12} = n(SSR_1 - SSR_2^{\min}) / SSR_2^{\min}$

where *n* is the observations associated with the values of z_{t-1} that lies between 15th and 85th percentiles (i.e., $n = 0.7(T-1)$). The *F* statistic has a non-standard asymptotic distribution under the SETAR hypothesis, so conventional critical values are not appropriate. Hansen (1999) showed how to obtain the appropriate critical value F_{12}^* using a bootstrapping procedure. The method involves resampling the data utilizing the residuals obtained from the above threshold model and for each bootstrap sample, searching the optimal threshold as we did before and calculating the test statistic F_{12}^*. This is repeated a large number of times (1000 in our case) to find the bootstrap distribution and thus the p-value for that representing the percentage of test statistics for which the test taken from the estimation sample exceeds the observed test statistics. This method will be applied to the full sample residuals obtained from the STC regression.

Before we proceed with the error correction procedures, two issues are worth discussing. First, we are investigating an adjustment process from a state-dependent two-regime nonlinear equilibrium, as opposed to most studies which analyze adjustment mechanisms on the basis of a one-regime linear equilibrium model. The dynamic adjustment mechanism between two different regimes does not have to be the same. Instead, it is plausible and reasonable that the error-correction process varies according to the "state" of equilibrium. For one thing, when world prices are unusually high, triggering active interventions; the adjustment pace for the domestic price to go back to the "active interventions" equilibrium therefore might be faster than it would be in a free market. This is especially true when dealing with less advanced economies because of imperfect information, high transaction costs, less developed infrastructure, restricted arbitrage, among many other institutional and economic conditions. To put this into consideration, we also investigate the error-correction processes using the subsample residuals split by long-run regime. We divide residuals into two groups according to the threshold value from STC estimation and then investigate the error correction processes under each regime correspondingly.

Another issue is that, it might be inappropriate to use a very short time period as a unit of reaction time span when investigating the error correction procedures. The model identification should reflect the reality that market reactions and adjustments may occur with a lag, especially for a transition economy. We therefore also consider multi-week differentials as a unit change in the "first-order difference" identification. That is, we identify the first-order of the error term as $\Delta_k z_t = z_t - z_{t-k}$ and its corresponding short-run response $\Delta_k y_t = y_t - y_{t-k}$, $k = 1, 2, ..., k_{\max}$ where k is the number of weeks that define a unit change, with k=1 as the special case usually applied in the literature. We then use the same SETAR method to test linearity utilizing the following equation

(20) $z_t = \alpha_1 z_{t-k} + e_t$ versus $z_t = \alpha_1 z_{t-k} I_1(\gamma) + \alpha_2 z_{t-k} I_2(\gamma) + e_t$.

We test linearity for three groups of residuals using different k values: the full sample residuals from STC, subsample residuals from STC regime 1 (world price below the threshold value), and subsample residuals from STC regime 2 (world price beyond the threshold value), with $k_{\max}$ =4. When we estimate models using k greater than one as a unit change, some observations are lost. To accurately compare the alternative models with different k value, the sample time period should be kept fixed (at T-$k_{\max}$ -lags). Otherwise, we would be comparing the performance of the models over different sample periods. The results are presented in Table 6. Model selection is based on AIC and SBC.

Table 6 about here

For all three groups of the sample, the Hansen tests do not reject the linearity hypothesis for all values of k. We then estimate the corresponding linear error correction models $\Delta_k y_t = \sum_{i=1} \alpha_i \Delta y_{t-i} + \sum_{j=1} \beta_j \Delta x_{t-j} + \lambda z_{t-k} + \varepsilon_t$ with k from 1 to 4 for all groups of residuals. Both AIC and SBC indicate that for each group of residuals, the case k=1 fits the best. We thus conclude the domestic price does respond to a deviation in a short time period. But as we will see, domestic price adjustments under both open trade and the active intervention regimes also respond to lagged price changes.

The results of error correction models when k=1 are presented in Table 7. We exclude the statistically insignificant regressors. First, for the full sample residuals, the adjustment of Ukrainian domestic price responds to the deviation from equilibrium and the lagged own price shocks and the world market price shocks. The results suggest that short-run dynamics of the Ukrainian prices react to the shocks from the world market with a lag of two and three weeks, but do not respond to shocks that occurred in the prior week. This was expected for an economy like Ukraine which has less developed market infrastructure and potentially high adjustment costs. To provide a little more intuition on the adjustment processes, we present the deviation half-lives for each group in Table 7.[vii] Adjustment towards the long-run equilibrium—takes place through changes in Ukrainian domestic wheat price alone—with half of the deviation from the equilibrium being corrected requiring nearly 18 weeks. The slow adjustment speed again may be a reflection of the institutional and economic characteristics of Ukrainian grain markets.

Table 7 about here

Next, we look at the subsamples. Under the open trade regime (regime 1), the price adjustment pattern is quite similar to the full sample situation. Price changes respond to disequilibrium and three-week lagged world price shocks, but not to the one- or two-week lagged changes. It takes roughly 18 weeks to eliminate half of the deviation from equilibrium, if changes occur only through the domestic price. In contrast, the adjustments under regime 2 are much faster. It costs only about eight weeks to eliminate half of the deviation from equilibrium. This is consistent with the fact that Ukrainian government always responded quickly and immediately to the rise in world grain prices over the sample period. For instance, in October 2006—right before the price crisis in 2007/2008, the Ukrainian government introduced a quote system as the world wheat prices start to increase. The quota volume was set at 0.4 million tons. Later in December 2006, the government dramatically reduced the quota volume, from 0.4 to 0.003 million tons (almost completely banned the wheat export) as a reaction to the continuous increases in world food price. The instrument it uses—a quota system—also makes the control take place quickly and effectively (from the view of controlling exports, not of improving the economy). Another interesting point is that adjustments of domestic price under this regime only respond to the deviations and to its own lagged price changes. It doesn't respond to the changes of lagged world market prices.

On the other hand, Ukraine is a major grain exporter. With intense world competition for commodities such as wheat, there is a legitimate concern that Ukraine may have some control over world market prices, at least in the short run. Some researchers and policy makers suggest that the export control in Ukraine is not only harming domestic markets and producers, but is also creating negative impacts on world grain markets and thus exacerbating the food crisis. We thus investigate whether world market prices also respond to deviations. We simultaneously estimate the error correction models for domestic and world prices using a seemingly unrelated regression technique. The results indicate that both under the full sample and the subsample situations, the world price does not respond to disequilibrium between the two markets. We also find that lagged changes in Ukrainian prices have no effect on adjustments of the world price. The results thus indicate that adjustments toward the long-run equilibrium take place through changes in Ukrainian prices alone. The result is consistent with the idea that the world market is large relative to Ukraine. This is also consistent with the 2008 World Bank report suggesting that although, export restrictions of grains and oilseeds/vegetable oils, of countries like Ukraine and Russia, have temporarily contributed to record world market prices. However, Ukraine's market power alone is limited in the long run and Ukraine would be ill-advised to attempt to exercise this influence by deliberately reducing exports in the long run in an effort to drive up world market prices and thus export revenues. However, our finding should not be interpreted as evidence that Ukraine has absolutely no effect on the world market price, but price shocks in Ukrainian domestic markets alone do not push the world market prices to make adjustments accordingly. Further investigation of the influence on the supply side would be helpful to understand the effects of Ukrainian trade interventions on world grain markets.

7 Conclusions

Our model results show that the wheat export restrictions in Ukraine during the two recent commodity price booms have temporarily changed the relationship between the domestic wheat grower price and the wheat world market price. To capture the effects of the export restrictions fully, a price transmission model with additional flexibility which allows for regime changes not only in the short-run but also the long-run price transmission parameters is required. This supports the findings of previous studies on price insulating trade policies (Myers and Jayne (2011) and Götz et al. (2010)).

In particular, our results indicate that the long-run price transmission decreased by 30% and the short-run speed of adjustment increased by 125%, compared to the open trade regime. The first finding is in line with our hypotheses on the grounds of the theory of the Walrasian equilibrium and the spatial price equilibrium theory. Though, several experts report substantial domestic wheat warehousing by traders in Ukraine in times of the export restrictions, which increased the domestic price and weakened the Walrasian equilibrium effect of the export quota. Thus, if additional wheat warehousing were not observed, the domestic price level would have been reduced more and the dampening effect on the magnitude of price transmission would have been even stronger. As explained above, almost no imports were observed during the export restrictions, which can be explained by a domestic price level lower than the world market price. Thus, the Walrasian equilibrium effect was not weakened by imports from the world wheat market.

Since the theory of spatial price equilibrium suggests that the speed of adjustment would have decreased as well, we follow that our second finding, the increased speed of adjustment in the closed trade regime is not induced by the export quota. Rather, we trace this back to the unusual large upward and downward price changes on the world market. We conjecture that large price movements cause increased price information flows and heightened attention to this information, which carries price changes from the world to the domestic market more efficient than physical trade flows. This information effect is rather strong and outweighs the dampening spatial price equilibrium effect such that the speed of adjustment increases.

The parameter estimates suggest that the Ukrainian wheat export quota has stabilized the domestic wheat price almost 30% below the international wheat price on average. However, to judge if the export quotas have been successful in insulating domestic wheat prices in Ukraine from world market prices, the effects of export restrictions not only on the domestic price but also on world market prices have to be taken into account. Martin and Anderson (2012) assess that almost 30% (for comparison: Anderson and Nelgen (2012a) estimate 19%) of the increase in world wheat prices 2006-2008 was caused by price insulating behaviour, i.e. the increase in export barriers by exporters as well as the removal of import barriers by wheat importing countries worldwide. Thus, our results indicate that the dampening effect of the export quota on the domestic price in Ukraine was fully compensated by the increasing effect of the changes in border protection rates on the world market price. This result is in line with the finding of Anderson and Nelgen (2012b) that governments were not successful in stabilizing domestic prices. From a global point of view, the domestic wheat price in Ukraine would have increased similarly, if no country had engaged in price insulating behaviour 2006-2008 worldwide – without causing any additional welfare costs. The welfare costs caused by the export quotas for Ukraine itself result from the foregone exports, the reduced wheat grower price, costs of the loaded ships locked in the harbour and the losses resulting from delayed or reduced investments in the Ukrainian grain production sector.

Currently, the implementation of export restrictions in the context of food security issues is in accordance with the WTO. In light of of the high welfare costs and the countervailing effect of trade insulating behaviour of exporting and importing countries world-wide, a new WTO law should be created which prohibits trade-policy interventions as a mean to improve food security. Alternatively, governments could help their needy people to cope with higher food prices by consumer-oriented measures, as e.g. direct income transfers or food vouchers.

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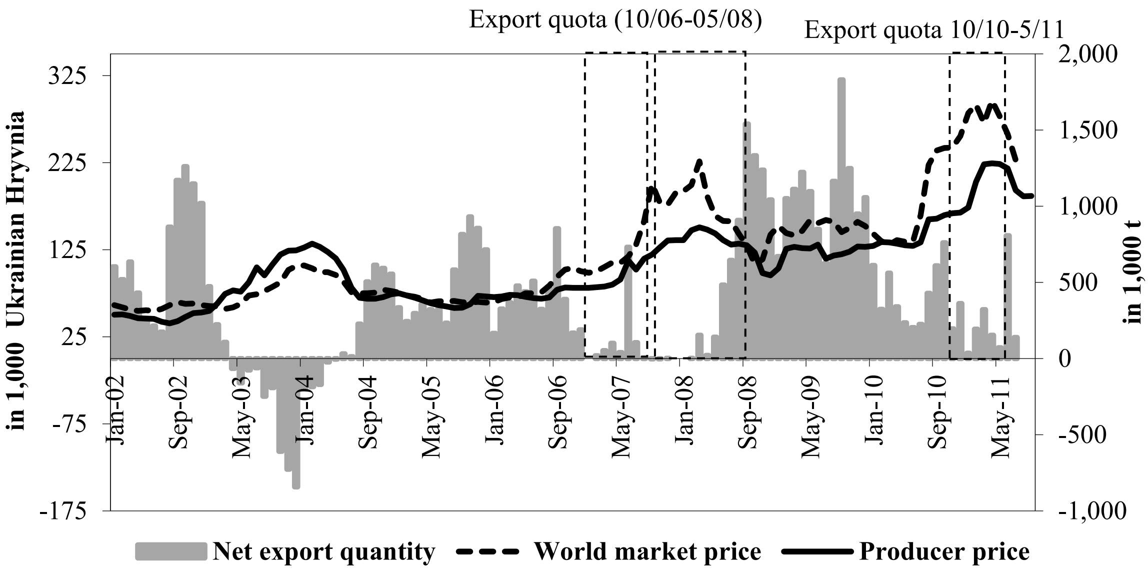
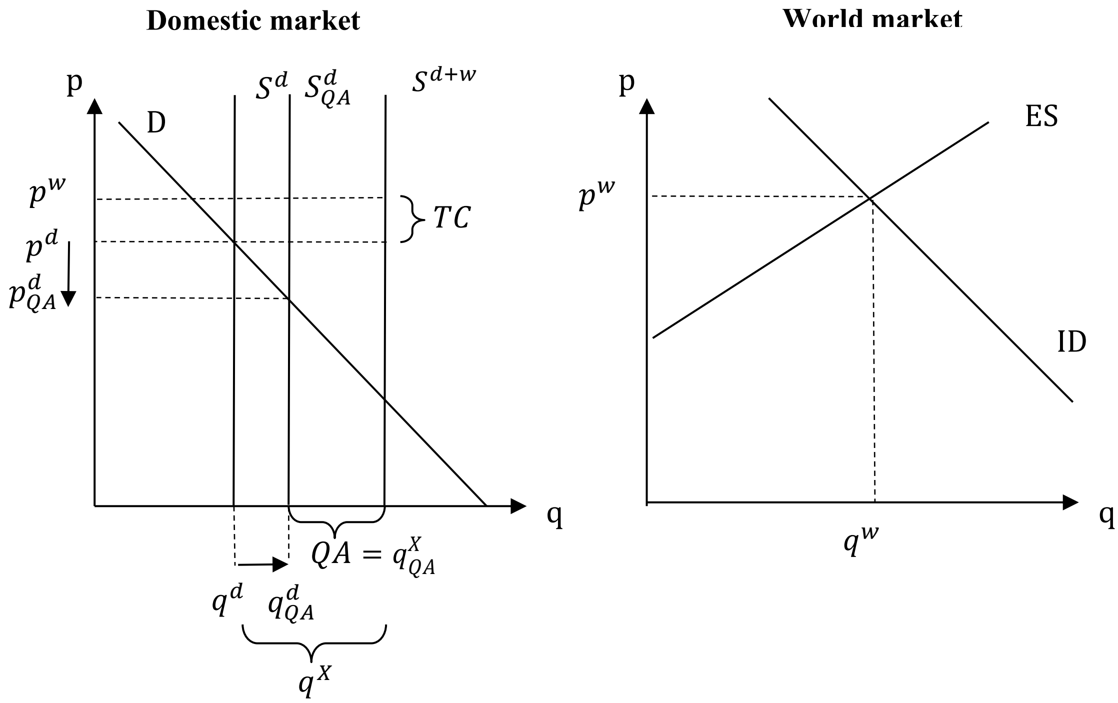
Figure 1: Development prices and exports Ukraine



Source: Own illustration; data base: GTIS (2011), APK-Inform (2011).

Figure 2: Price and Quantity Effects of an Export Quota on the Domestic Market



Source: Own illustration.

Figure 3: Ukraine and world market prices (US$/ton): March 16, 2001-September 9, 2011

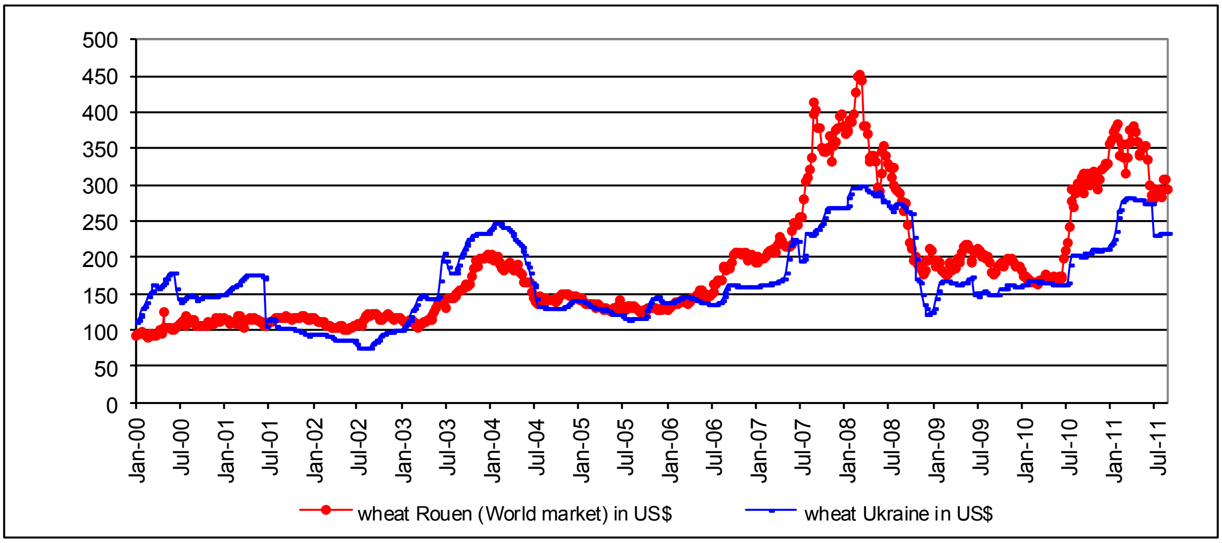
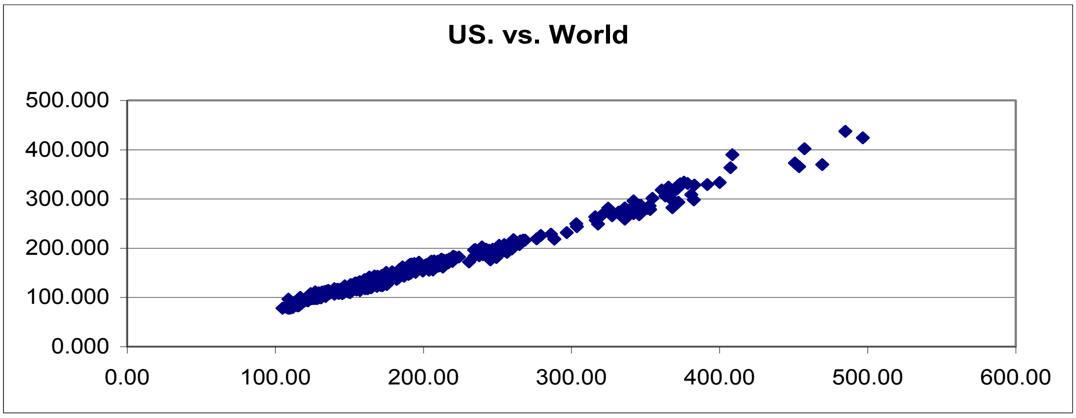
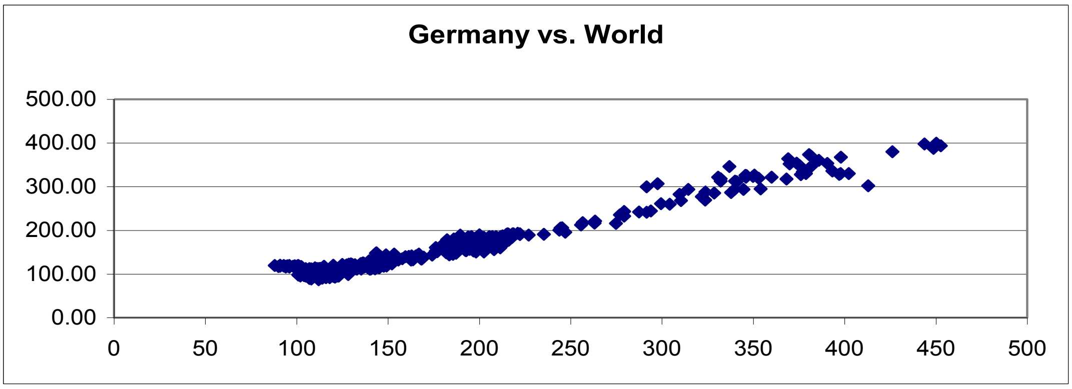


Figure 4: Domestic and its corresponding world market prices.

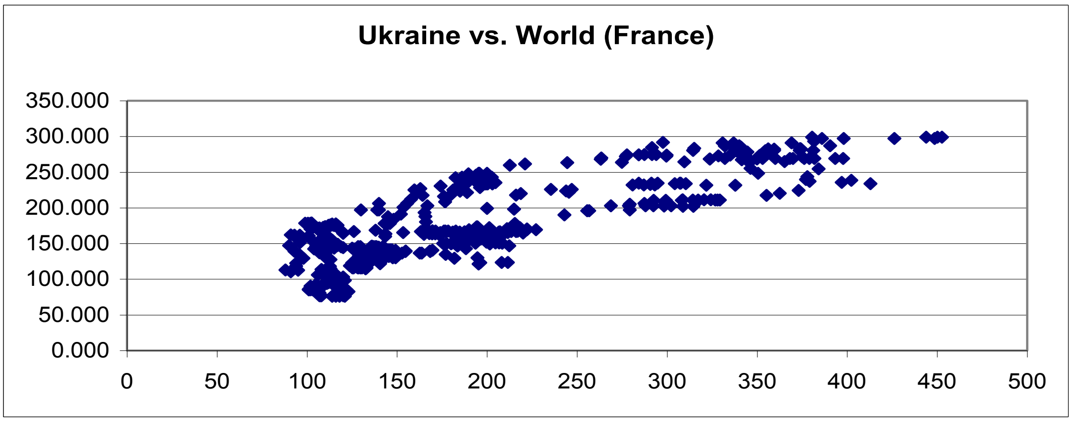
a)



b)



c)



Note: For the U.S., the FOB price of hard red winter wheat at the USA Gulf port (HGCA 2009) has been utilized as the relevant world market price for the USA; and for the Germany and Ukraine, the world reference price is the FOB price of wheat (classification "other wheats") in Rouen, France (HGCA 2009).

Figure 5: Smooth Transition Cointegration Model Fit

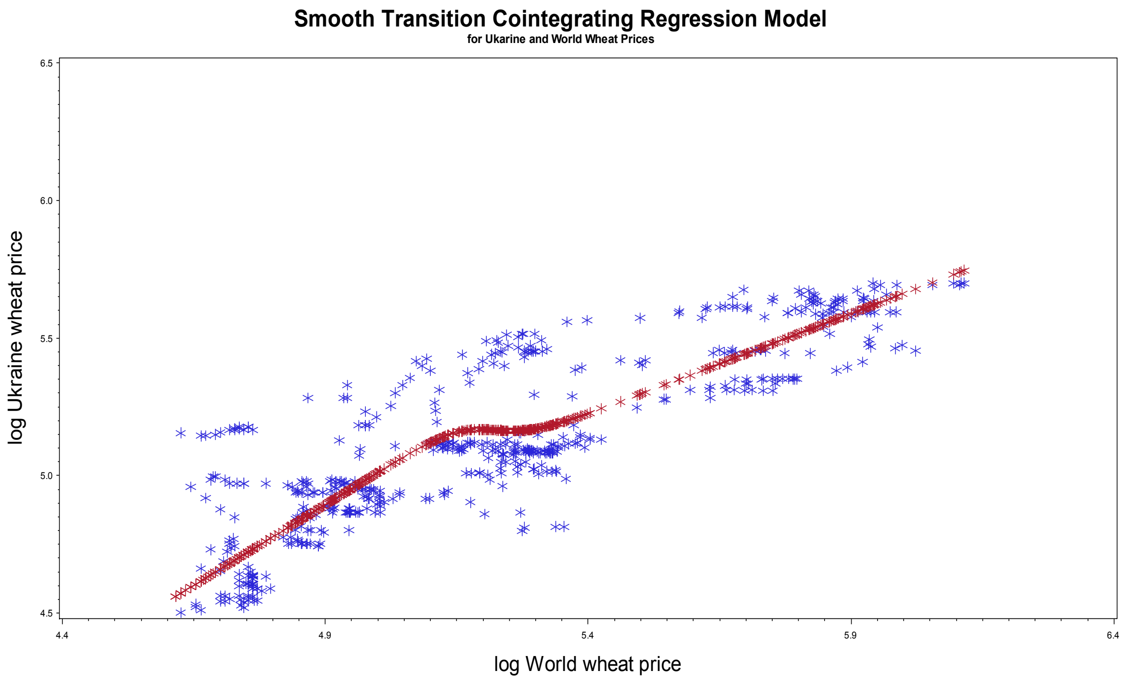


Table 1: Theoretical Effects of Export Restrictions on the Relationship between the Domestic and the World Market Price

	Magnitude of price transmission	**Speed of price transmission**
1) Physical trade effect	↓	-
2) Domestic supply effect	↓	↓
3) Information effect	-	↑

Source: Own illustration.

Table 2: Unit Root Tests for Price Data (in natural logarithms)

		World price		Ukraine price		Δ World price		Δ Ukraine price
Dickey-Fuller								
Single Mean	Lags		Lags		Lags		Lags	
ρ	3	-3.15	3	-6.81	3	-319.74	3	-252.802
(Pr < ρ),		(0.64)		(0.29)		(<0.001)		(<0.001)
τ_{μ}	3	-1.19	3	-1.75	3	-9.95	3	-9.23
(Pr < τ_{μ})		(0.68)		(0.40)		(<0.001)		(<0.001)
Trend								
ρ	6	-8.54	3	-12.98	6	-319.78	6	-8.54
(Pr < ρ),		(0.54)		(0.26)		(<0.001)		(<0.001)
τ_{μ}	6	-2.05	3	-2.66	6	-9.94	6	-253.78
(Pr < τ_{μ})		(0.58)		(0.25)		(<0.001)		(<0.001)
KPSS								
Single Mean	6	4.81	6	2.93	6	0.07	6	0.10
Trend	6	0.30	6	0.26	6	0.07	6	0.08

Note: The 10%, 5%, and 1% critical values for KPSS-single mean test are 0.35, 0.46, and 0.74, respectively; and for KPSS-trend test are 0.12, 0.15, and 0.22 respectively.

Table 3: Linear vs. Smooth Transition Cointegrating Vector Tests

	Ukrainian vs. world market price	U.S. vs. world market price	German vs. world market price
Lags and Leads $\sum_{j=-K}^{K} \alpha_j \Delta p_j^{wd}$		Statistic τ (3rd order Taylor approx.)	
K=1	12.83	0.88	1.13
K=2	11.99	0.39	1.05
K=3	12.17	0.54	0.87

The tau statistic follows a chi-square distribution with two degree of freedom. The null hypothesis is linear cointegrating vector and the alternative is STC. The critical value is $\chi(2)_{0.05} = 5.99$.

Table 4: Estimates of the Smooth Transition Cointegrating Models

	STC, no lags and leads			STC, with lags and leads		
Parameter	Estimate	Approx Std Err	Approx Pr > |t|	Estimate	Approx Std Err	Approx Pr > |t|
γ	3.87	1.73	0.03	3.23	1.18	<0.01
c	5.21 ($185)	0.05	<0.01	5.17	0.05	<0.01
α_1	-0.86	0.49	0.08	-1.45	0.50	<0.01
α_2	2.13	0.67	<0.01	2.77	0.69	<0.01
β_1	1.14	0.10	<0.01	1.19	0.10	<0.01
β_2	-0.44	0.13	<0.01	-0.57	0.13	<0.01
π_{t+1}^0				-0.44	0.45	0.34
π_{t+1}^1				0.59	0.54	0.27
π_t^0				-0.48	0.45	0.29
π_t^1				-0.19	0.54	0.73
π_{t-1}^0				-0.22	0.46	0.64
π_{t-1}^1				-0.54	0.56	0.33
$\sum (y_t - \hat{y}_t)^2$	8.21			7.54		

Table 5: Stationarity Tests for Residuals Obtained from estimated STC model

	Lags	Statistics
Engle-Granger Cointegration Test	3	-32.88
KPSS, Single Mean	6	0.41

Note: The 10%, 5%, and 1% critical values for KPSS-single mean test are 0.35, 0.46, and 0.74, respectively; and for KPSS-trend test are 0.12, 0.15, and 0.22 respectively. The 10%, 5%, and 1% critical values for Engle-Granger cointegration test (with two variables, sample size 500, and include a constant in the cointegrating vector) are -3.05, -3.35, and -3.92, respectively.

Table 6: Residual-based Tests of Linearity, Hansen F-test

	Bootstrap P-value for Hansen 1999 test		
	Full sample residuals	Residuals from STC regime 1 (world price <= $185)	Residuals from STC regime 2 (world price > $185)
$\Delta_k z_t = z_t - z_{t-k}$			
k=1	0.93	0.49	0.43
k=2	0.90	0.45	0.42
k=3	0.90	0.44	0.44
k=4	0.92	0.46	0.42

Table 7: Estimates for Linear ECMs

	All residuals		Residuals from regime 1		Residuals from regime 2	
Variable	Coef.	Std Err	Coef.	Std Err	Coef.	Std Err
z_{t-1}	-0.04	0.009	-0.04	0.014	-0.09	0.018
Δy_{t-1}	0.23	0.042	0.10	0.058	0.19	0.061
Δy_{t-2}	0.21	0.066				
Δy_{t-3}	0.21	0.067	0.33	0.107	0.13	0.063
Δy_{t-4}						
Δx_{t-1}						
Δx_{t-2}	-0.14	0.053				
Δx_{t-3}	-0.13	0.054	-0.20	0.095		
Δx_{t-4}						
Half-life	17.7wks		17.7wks		8.0wks	
AIC	-294.92		-207.13		-193.50	
SBC	-252.42		-174.20		-182.99	
Observation	542		250		292	

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[i] We also test unit root using a nonparametric, residual-based stationary bootstrap test developed by Parker, Paparoditis, and Politis 2006 (PPP thereafter). The PPP procedure offers significant improvements over the large sample Gaussian approximations commonly used in the econometric analysis of non-stationary time series, as it does not rely on a specific data generating process. The test results are consistent with ADF results.

[ii] If not, then follow the common practice and estimate the linear cointegration.

[iii] Choi and Saikkonen (2004) argue that because the motivation for using the third-order instead of the first-order approximation is to improve the power of test statistics, they thus suggest using third-order approximation only for the transition of the intercept term and using the first-order approximation for the transition involving the regressors.

[iv] Many empirical studies may utilize maximum likelihood methods in application. Under the additional assumption that the errors of Equation (5) are normally distributed, NLS is equivalent to maximum likelihood. Otherwise, the NLS estimates can be interpreted as quasi-maximum likelihood estimates.

[v] We also test the unit root and stationarity of the residuals by regime. In particular, we split the residuals into two groups by regime and conduct the ADF, bootstrapping, and KPSS tests accordingly. The results are consistent with the test results obtained from the full sample residuals.

[vi] Due to severe winter-kill, the smallest harvest in more than 45 years was produced in marketing year (MY) 2003/2004 in Ukraine, which made Ukraine a wheat importer in that year. This one exception a, side, Ukraine is a net wheat exporter over our sample time period.

[vii] Deviation half-lives, given by $\ln(0.5)/\ln(1+\lambda)*k$, where λ is the OLS estimate of $\Delta z_t = \lambda z_{t-1} + \varepsilon_t$ or $\Delta y_t = \sum \alpha_i \Delta x_i + \sum \alpha_i \Delta y_i + \lambda z_{t-1} + \varepsilon_t$, represent the period of time (in weeks) required for one-half of a deviation from equilibrium to be eliminated.