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EXPORT RESTRICTIONS AND MULTIPLE SPATIAL PRICE EQUILIBRIA: EXPORT QUOTAS FOR WHEAT IN UKRAINE

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EXPORT RESTRICTIONS AND MULTIPLE SPATIAL PRICE EQUILIBRIA: EXPORT QUOTAS FOR WHEAT IN UKRAINE

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Abstract

Relatively few models exist that allow for regime-dependent spatial price equilibria. This paper focuses on temporary export restrictions during international commodity price peaks. Theory suggests that export restrictions have price insulating effects and lead to multiple spatial equilibria between domestic and world market prices. Our analysis is unique in that it tests for linear versus non-linear cointegration within a smooth transition cointegration model. Applying this model to the wheat export quota in Ukraine shows that the domestic wheat price was stabilised approximately 30% below the international wheat price during the two recent price spikes. 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 worldwide from 2006 to 2008.

Keywords: Agricultural policy, Food crisis, Price transmission, export controls, Ukraine

1 Introduction

During the recent price booms on world agricultural markets in 2007-2008 and 2010-2011, many countries aimed to insulate their domestic markets from price developments on the world market and to stabilise domestic prices through trade policy interventions. Exporting countries implemented export restrictions reducing or even banning exports, and importing countries reduced or even completely eliminated import restrictions to dampen the influence of high world market prices on domestic price levels. Most studies on trade policy interventions in the context of commodity price peaks focus on world market price effects (e.g. Martin and Anderson, 2012; Anderson and Nelgen 2012a); there are relatively few studies that investigate the influences that these interventions have on domestic prices (e.g. Anderson and Nelgen 2012b; Götz et al. 2013; Grueninger and von Cramon-Taubadel 2008).

This paper adopts an empirical framework to identify and measure the effects of export restrictions on the relationship between domestic and world market prices. The question naturally arises over how successful export restrictions were in insulating the domestic price from the world market price. Theory suggests that export restrictions reduce the magnitude of price transmission from the world market to the domestic market. To capture these effects, we develop a flexible spatial price transmission model that allows for a regime-dependent long-run price equilibrium relationship. We apply this model to the Ukrainian wheat market, where exports were restricted by an export quota system during price peaks in 2007-2008 and 2010-2011.

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Recent methodological innovations in spatial price transmission analysis have led to increased model flexibility. For example, the threshold vector error correction model (TVECM) and the threshold autoregressive (TAR) model (Balke and Fomby 1997) are motivated by the existence of transaction costs (Goodwin and Piggott 2001). As long as the deviations from the equilibrium are smaller than the transaction costs, which are represented by a threshold in the model, the "neutral band" regime with insignificant or low adjustment speed prevails. When the deviations from the equilibrium exceed the threshold, an "outside-band" regime characterised by significant error correction behaviour prevails. The smooth transition vector error correction model (STVECM) or the 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 the abrupt regime change in the TVECM and TAR models. This more flexible approach allows for inhomogeneity or geographically dispersed agents, both of which may lead to differing transaction costs (Goodwin et al. 2011).

Although the abovementioned 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 the magnitude in price transmission is linear and constant. However, multiple long-run price equilibria might exist as well with changing long-run price transmission parameters. Relatively few applications of spatial price transmission models allow for regime-dependent long-run spatial price equilibria. 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. 2013) and the composition of trade flows (Götz et al. 2008).

Our spatial price transmission analysis is unique in that it introduces a new test for linear versus non-linear cointegrating relationship developed by Choi and Saikkonen (2004). We estimate the long-run price transmission parameters in the framework of a smooth transition cointegration model based on Saikkonen and Choi (2004), and allow for a two-regime long-run price transmission relationship, with the world market price being the variable that determines transition from one regime to the other.

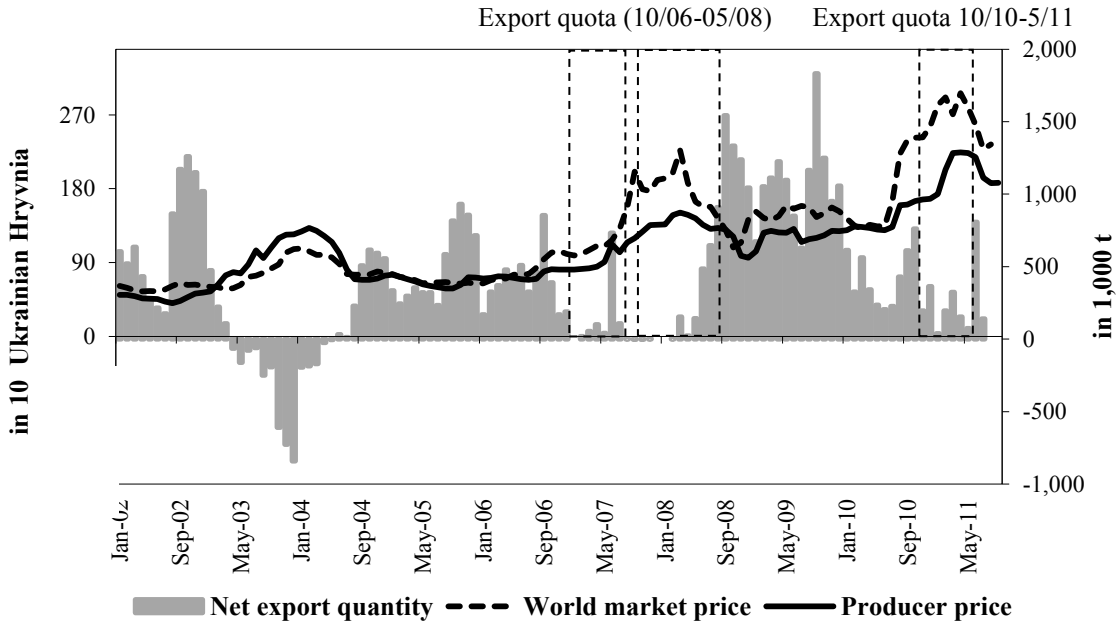
The remainder of this paper is structured as follows. Section 2 presents background information on the Ukrainian wheat export quota. Section 3 reviews the most relevant articles and explains differences regarding our approach. The data base and the empirical approach are explained in section 4 and estimation results are presented in section 5. Conclusions are drawn in section 6.

2 The Ukrainian Wheat Export Quota

The Ukrainian government quantitatively limited wheat exports during the two recent commodity price spikes through an export quota that was implemented within a governmental licensing system. Export quotas allow exports up to the amount 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) along with net wheat exports. Ukraine became a net wheat importer from 2003-2004 when the Ukrainian price increased above the world market price; however, wheat imports were so low that they were almost not observed when the export quota system was effective. During periods of

export restrictions, the domestic wheat price was below the world wheat market price, and the difference between these two prices increased.

Figure 1: Development Prices and Exports, Ukraine

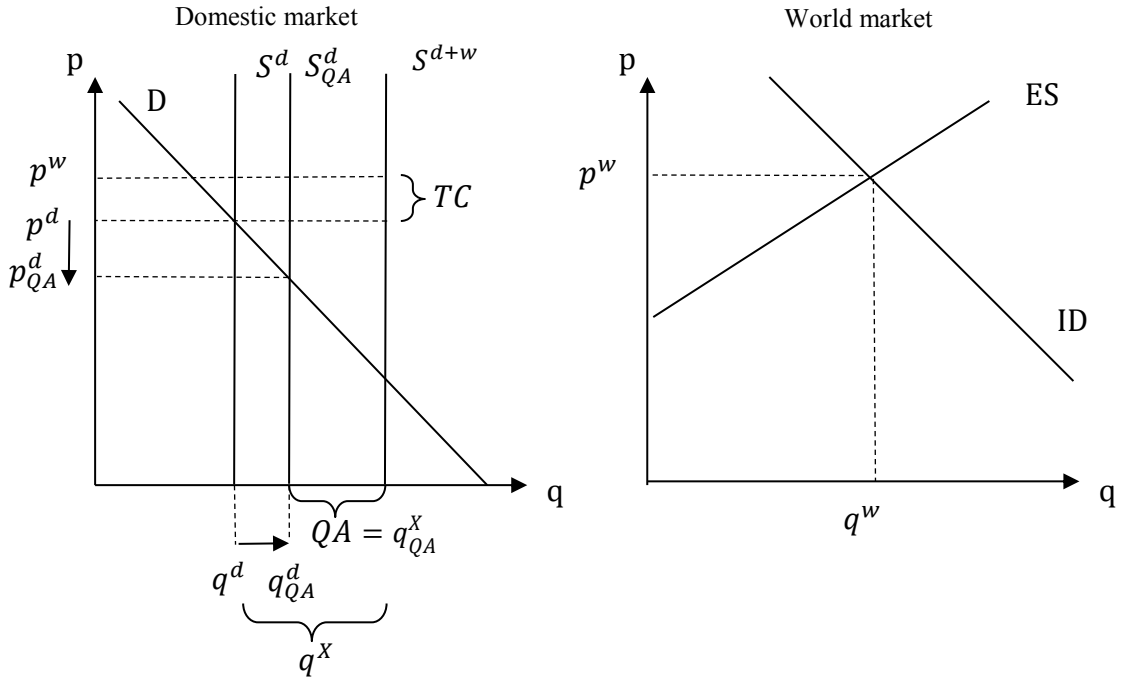


Source: Own illustration. Data base: GTIS (2011), APK-Inform (2011).

This price development can be explained by the theory of the Walrasian equilibrium, which clarifies how export restrictions change supply and price on the domestic market. In particular, the size of exports decreases and more is supplied to the domestic market, which reduces the domestic market price compared to the price that would prevail if trade was possible. 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. Implementing the export restriction reduces the export quantity q^x to $q^d_{QA} = QA$, the amount 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 that 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 the domestic price decreases. 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 Ukrainian wheat export quota system was 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 its distribution suffered from massive corruption, particularly in 2010-2011. For example, the wheat export quota implemented in 2010 became effective so rapidly 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, thereby causing high additional costs to exporters (APK Inform 2010). According to traders' information, this implied that contracts could not be fulfilled, which negatively affected the international reputation of traders exporting from Ukraine.

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



Source: Own illustration.

3 Literature Review and Model Selection

The effects of wheat export restrictions on the integration of the Ukrainian market into world markets are also investigated by Götz et al. (2013). These authors use a Markov-switching vector Error-correction model (MSVECM) assuming that the regime switches in the MSVECM (Krolzig et al. 2002) are governed by an unobserved and probabilistic variable. Focusing on the wheat export controls in Russia and Ukraine during the 2007-2008 food crisis, three price transmission regimes are identified for Russia and Ukraine, with the "crisis" regime prevailing mainly during times of export restrictions. This "crisis" regime is characterised by reduced long-run price transmission.

The main difference between our smooth transition model approach and the MSVECM framework in Götz et al. (2013) is that the regime switches in the long-run equilibrium regression are not abrupt, but are assumed to 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 supposed to be determined by the world wheat market price. We assume that although wheat export controls in Ukraine were implemented abruptly, some traders might have already reacted before their implementation. When traders expect that world market prices will further increase and rumours circulate that export restrictions might be implemented, traders might already change their behaviour. In particular, some traders might increase their export activities temporarily until the exports become restricted or forbidden. Other traders might reduce their export activities to prevent potential losses if they expect exports will be restricted or even banned on short notice (compare section 2). Since the point in 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 us to explicitly test for non-linearity in the long-

run cointegration regression. In the MSVECM framework in Götz et al. (2010), a Likelihood-ratio test for general non-linearity of the model is conducted. Goychuk and Meyers (2011) investigate the integration of wheat markets in Russia and Ukraine with world wheat markets in 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 in a linear model approach. Ghoshray (2010) analyses the cointegration of international wheat prices (USA, Argentina, Australia, Canada and the EU) 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 a smooth rather than an abrupt change between the "outside-band" and the "neutral band" is motivated by the potential variability of transaction costs influenced by trade volumes. Cointegration is confirmed for all price pairs, indicating highly integrated world wheat markets. The main difference between the model approach in Ghoshray (2010) and our smooth transition model approach is that the long-run equilibrium relationship is assumed to be linear; instead the behaviour of the speed of adjustment is assumed to be non-linear, depending on the size of the deviation from the equilibrium. Gervais (2011) captures nonlinearity in the long-run and short-run equilibrium relationships in an investigation of 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, but results do not indicate asymmetry in the speed of price transmission. In contrast, 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. Relatively few studies in spatial price transmission exist that allow for non-linearity in the long-run price transmission. One example is Stephens et al. (2012), who estimate a regime-specific cointegration model for tomato markets in Zimbabwe by distinguishing between two price transmission regimes in the presence and absence of trade. Their model specification allows the two regimes to differ in both short-run and long-run price transmission. They also identify non-linear cointegration and observe error correction behaviour in the absence of trade. Still, long-run price transmission and the speed of adjustment are both higher in the no-trade regime compared to the trade regime. Myers and Jayne (2011) investigate spatial maize price transmission between South Africa and Zambia, and propose a multiple-regime threshold model with changing price transmission regimes, allowing for multiple speed of adjustment as well as multiple long-run equilibria. The regime switches are assumed to depend on the magnitude of trade flows between the regions, temporary governmental market interventions and whether transport capacity constraints are binding. When Zambian imports are high, the government is the main importer and sells maize at subsidised prices, which reduces the domestic market price. This implies that cointegration is not confirmed and a long-run price equilibrium does not exist. Götz and von Cramon-Taubadel (2008) develop a procedure to estimate a regime-dependent VECM, which allows both the short-run and the long-run price transmission to differ between regimes. 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, for apple price data from two German wholesale markets. The test identifies four price transmission regimes that differ in the equilibrium relationships.

4 Empirical Strategies

To capture the effects of temporary export restrictions on the integration of the Ukrainian domestic wheat market and the world wheat market, we choose the smooth transition cointegrating (STC) framework of Saikkonen and Choi (2004) and Saikkonen and Choi (2004), and follow the general procedure suggested by Engle and Granger (1987). The empirical procedure used to analyse the regime-switching price transmission can be broken down into the following steps: 1) Test for a 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 favour of STC (as in our case), using the method proposed by Saikkonen and Choi (2004);[3] 3) Test stationarity using the residuals obtained from the estimated STC model; 4) Test linearity versus nonlinearity of the error correction procedure, based on the residuals of the estimated STC regression model; 5) Estimate the error correction model, based on the test results from 4), to investigate the dynamic adjustments between the Ukrainian and the world market wheat price.

Test on Linear versus STC Long-run Relationship

Testing linearity against the STC specification constitutes a first step towards building the STC model. We adopt a test based on Choi and Saikkonen (2004). Consider a general STC model:

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

where y_t denotes the dependent variable and x_t represents the *I(1)* independent variable(s); 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 transmission elasticity under the price transmission framework, and $g(x_t - c; \gamma)$ is a smooth transition function of the process x_t, with a 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 considered a regime-switching model that allows for two regimes associated with extreme values of the transition function, i.e. $g(x_t - c; \gamma) = 1$ 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[4] give rise to different types of regime-switching behaviour. In our study we use a first-order logistic function as the transition:

$$g(x_t - c; \gamma) = \left[1 + \exp(-\gamma(x_t - c))\right]^{-1}. \qquad (2)$$

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 behaviour 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 behaviour 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

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

[4] For detailed discussions on the choice of transition functions, the reader is referred to van Dijk, Teräsvirta, and Franses (2002), and Teräsvirta, Tjøstheim, and Granger (2010).

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*). In our study, wheat's world market price serves as the transition variable x_t that determines the switch between the two regimes. We assume that if the world market price exceeds a certain level, i.e. the threshold value c, the Ukrainian government imposes temporary export restrictions that lead to a change in the relationship between the domestic Ukrainian and wheat's world market price. Due to heterogeneity of the traders (compare section 3), the transition process is smooth. We expect that the long-run price transmission regime prevailing when trade is open will be different from the regime in times of restricted wheat exports.

The null hypothesis of linearity can be expressed as equality of the autoregressive parameters in the two regimes of the STC model in (1). 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. Choi and Saikkonen (2004) develop a nonlinearity test that extends the approaches developed by Luukkonen et al. (1988) and Granger and Teräsvirta (1993), and can be applied in the context of STC. In particular, their test relaxes the exogeneity requirement for the regressors and follows the common practice of cointegrating regression and permits both serial and contemporaneous correlations between the regressors and the model's error term. 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 (1) is small. We thus adopt the third-order Taylor approximation and rewrite the transition function as:

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

The testing procedure involves estimating the corresponding auxiliary regression using OLS:[5]

$$\begin{aligned} y_t &= \alpha_1 + \alpha_2 \left\{ b\gamma(x_t - c) + d\left[\gamma(x_t - c)\right]^2 + h\left[\gamma(x_t - c)\right]^3 \right\} \\ &\quad + \beta_1 x_t + \beta_2 x_t b\gamma(x_t - c) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} \qquad (4) \\ &= \omega + \phi_1 x_t + \phi_2 x_t^{\ 2} + \phi_3 x_t^{\ 3} + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \quad t = K+1,\ldots,T-K \end{aligned}$$

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

[5] Choi and Saikkonen (2004) argue that 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.

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.

Estimation of the STC Long-run Relationship

Given that the null hypothesis of linearity has been rejected, the next step is to estimate the STC equation (1). We adopt the STC regression framework developed by Saikkonen and Choi (2004), where regressors are *I(1)* and errors are *I(0)*, and the regressors and errors are allowed to be dependent both serially and contemporaneously. Previous studies (e.g. van Dijk, Teräsvirta, and Franses 2002) usually suggest using a nonlinear least square (NLLS) technique to obtain the estimates of the parameters in (1). Saikkonen and Choi (2004) indicate that although the nonlinear least squares estimator is consistent, the asymptotic distribution involves a bias if regressors and error are dependent, which makes the NLLS estimator inefficient and unsuitable for hypothesis testing. They thus propose a Gauss–Newton (G-N) type estimator that utilises the NLLS estimator as an initial starting value and expand the model by including leads and lags as extra regressors. We therefore adopt their iterative estimation procedure and utilise a damped G-N method, known as the Levenberg-Marquardt (L-M) method. Given that the initial values of the parameters are close to the final optimal values, the L-M method has proven to be more efficient and can almost always guarantee quadratic final convergence. Also, as discussed, the results of our estimation could be sensitive to the initial values of γ and c. Thus, van Dijk, Teräsvirta, and Franses (2002) suggest normalising 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 (2) with a normalised version:

$$g(x_t - c;\gamma) = \left[1 + \exp(-\frac{\gamma}{\hat{\sigma}_x^{\,2}}(x_t - c))\right]^{-1}. \qquad (5)$$

Finally, an error correction model may be estimated as follows:

$$\Delta y_t = \alpha_0 + \sum_{m=1}^{M} \theta_{1m} * \Delta y_{t-m+1} + \sum_{l=1}^{L} \theta_{2l} * \Delta x_{t-l+1} + \delta * z_{t-1} + \omega_t, \qquad (6)$$

with t=1,2,…T and

$$z_{t-1} = y_t - (\alpha_1 + \beta_1 x_{t-1}) - (\alpha_2 + \beta_2 x_{t-1}) g(x_{t-1} - c;\gamma). \qquad (7)$$

5 Data and Results

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 (compare Figure 1).

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). 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 (Detailed test results are available from the authors upon request). 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 may be considered as I(1) processes.

Test Linear versus STC Long-run Relationship

The STC relationship test results are presented in Table 1. 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 model for the long-run relationship between the Ukrainian and world market wheat prices. As a comparison, we also test linearity of the relationship of German and US wheat prices with their corresponding world market prices. Neither of the tests is able to reject the linearity assumption. This is consistent with our priors, since these two countries did not implement trade restrictions during the food crisis. In the next step, we estimate the STC relationship for Ukraine. To decide if the Ukrainian and the world market prices are indeed cointegrated, and thus if a long-run equilibrium exists, we test stationarity of the residuals obtained from the above STC regression. 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 the unit root. We therefore conclude that the Ukrainian and world market wheat prices are cointegrated via a smooth transition mechanism.

Table 1: Test Results, Linear vs. Smooth Transition Cointegration

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

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

STC Regression Model

Table 2 presents the (iterated) L-M estimates of the cointegration model for the Ukrainian and world wheat markets. The STC results from Table 2 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 regressor-error dependence is not an issue in our sample set. Parameter estimates for the STC long-run equilibrium relationship are presented in (8), which are retrieved from the STC with no lags and leads:

$$\hat{y}_t = \begin{cases} -0.86 + 1.14x_t, & \text{if g=0} \\ 1.27 + 0.70x_t, & \text{if g=1} \end{cases}, \quad \text{and } g(x_t - c;\gamma) = 1/\{1 + \exp[-3.87(x_t - 5.21)/0.16)]\} \quad (8)$$

The results confirm a regime-switching behaviour in the long-run relationship between the Ukrainian domestic and world market prices, depending on the level of world market prices. The estimated threshold value for the transition variables is 5.21 in logarithms, or 185. Thus, when the world price is below the threshold value of $185/ton, the transmission elasticity of the domestic price with respect to the world price tends to be 1.1, suggesting that the two markets are closely integrated. This provides evidence that when the world price is not "too high" (below the threshold value of $185/ton), no active export controls are triggered, and price changes or shocks on the world market are fully transmitted to the Ukrainian market. Conversely, when the world market is "too high" (from the perspective of the Ukrainian government), and exceeds the threshold level of $185/ton, the relationship between the two markets gradually switches from the "open trade" regime to the "restricted trade" regime, which is characterized by a lower long-run transmission elasticity of 0.70.

Figure 3 shows the time periods when the "restricted trade" regime prevails according to our model results; the "open trade" regime prevails otherwise. It becomes evident that the time period when the export quota system applies (October 2006-May 2008; October 2010-May 2011) is fully captured by the "restricted trade regime", but is not congruent with it. When high world market prices prevailed from 2003-2004, Ukraine turned into a net wheat importer (see Figure 1); thus, export restrictions were not required. In addition, world market prices exceeded the threshold value only slightly for a short period of time. Export restrictions were also not implemented in times of relatively high world market prices in the marketing year 2008/09 since Ukraine experienced a bumper harvest in 2008, and supply on domestic markets was high.

Table 2: Estimates of the Smooth Transition Cointegrating Model

Parameter	Estimate	Approx. Std. Err.	Approx. Pr > |t|
γ	3.87	1.73	0.03
c	5.21 ($185)	0.05	<0.01
α_1	-0.86	0.49	0.08
α_2	2.13	0.67	<0.01
β_1	1.14	0.10	<0.01
β_2	-0.44	0.13	<0.01
$\sum(y_t - \hat{y}_t)^2$	8.21		

Finally, during the transition phase from the "open trade" regime to the "restricted trade" regime, an increase of one unit in the world market price was only partially passed along to the domestic market while a similar decrease in the world price was fully transmitted to the domestic market. Under such circumstances, the domestic growers are thus worse off from price increases compared to the potential benefit they might gain from the same price increase on the world market, all else being equal.

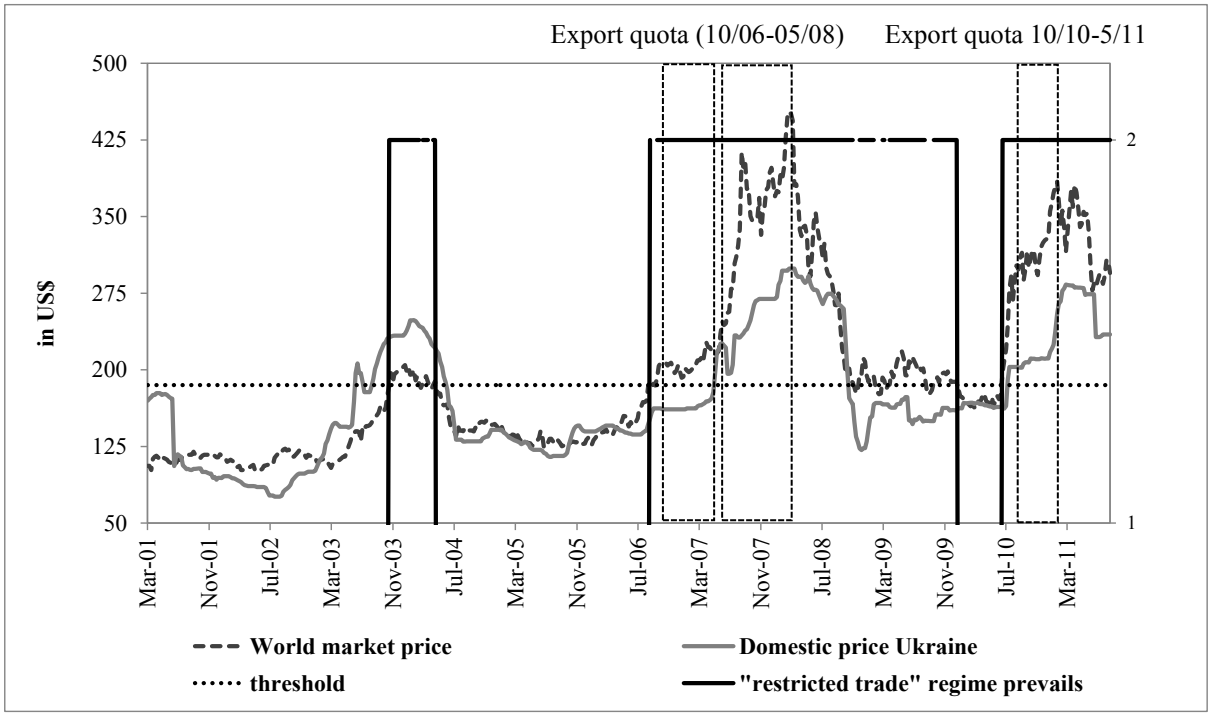
Estimate the Error Correction Model

The transaction cost version of the LOP provides justification for using a TAR or STAR model type, the momentum threshold autoregressive (M-TAR) or exponential STAR types of regime-switching models that allow the speed of adjustment behavior to be asymmetric inside

compared to outside the transaction cost band. We first conduct a linearity test for the residuals, which is based on Hansen's (1999) self-exciting threshold autoregressive (SETAR) approach. The bootstrapping p-value for the Hansen test is 0.93, thus the null hypothesis of linearity cannot be rejected (Table 3). Therefore, we estimate a linear error correction model in the next step.

The estimated parameters for the error correction model are presented in Table 4. Lag selection is based on the AIC and SBC, and we exclude the statistically insignificant regressors. Results indicate that the Ukrainian domestic price responds to the deviation from the equilibrium with the world market price, although at a rather low speed. The estimate indicates that 4% of the deviation is corrected in the following period. Also, the Ukrainian price changes significantly depending on the lagged own price shocks, as well as the lagged world market price shocks up to a lag of three weeks each. To provide more insight to the adjustment processes, we also present the deviation half-life in Table 5. The half-lives presented in Table 5 show that it requires nearly 18 weeks correcting fifty percent of a deviation from the equilibrium. The rather slow adjustment speed and high half-lives can be explained by the institutional and economic characteristics of Ukrainian grain markets. Grain traders confirm that the wheat export market in Ukraine is characterized by high marketing costs due to an underdeveloped market

Figure 3: **Regime Affiliation According to Model Results**



Note: regime 1= "open trade" regime; regime 2= "restricted trade" regime; whenever the "restricted trade" regime does not prevail, the "open trade" regime applies. Source: Own illustration.

6 Conclusions

The results of our smooth transition cointegration approach illustrate that wheat export restrictions in Ukraine during the two recent commodity price spikes have changed the relationship between the domestic wheat grower price and the wheat world market price. We captured these effects in a price transmission model with additional flexibility, which allows

for regime change in the magnitude of price transmission. This supports the findings of previous studies on price-insulating trade policies (Myers and Jayne 2011; Götz et al. 2013).

Our results particularly indicate that the long-run price transmission decreased by about 30% compared to the open trade regime. This finding is in line with our hypotheses based on the theory of the Walrasian equilibrium. However, several experts reported substantial domestic wheat warehousing by Ukrainian traders during the export restrictions, which increased the domestic price and weakened the effects of the binding export quota. Thus, if additional wheat warehousing were not observed, the domestic price level would have been further reduced 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 domestic prices being lower than world market prices. Thus, the effect of the export quota system on domestic supply was not weakened by imports from the world wheat market.

Table 3: Residual-based Tests of Linearity, Hansen F-test

Bootstrap P-value for Hansen 1999 test	
$\Delta_k z_t = z_t - z_{t-k}$	
k=1	0.93
k=2	0.90
k=3	0.90
k=4	0.92

Table 4: Estimates for Linear ECMs

	All residuals	
Variable	Coef.	Std Err
z_{t-1}	-0.04	0.009
Δy_{t-1}	0.23	0.042
Δy_{t-2}	0.21	0.066
Δy_{t-3}	0.21	0.067
Δy_{t-4}		
Δx_{t-1}		
Δx_{t-2}	-0.14	0.053
Δx_{t-3}	-0.13	0.054
Δx_{t-4}		
Half-life	17.7wks	
AIC	-294.92	
SBC	-252.42	
Observation	542	

The parameter estimates suggest that on average, the Ukrainian wheat export quota stabilised domestic wheat prices almost 30% below international wheat prices. However, to judge whether the export quotas were successful in insulating Ukrainian domestic wheat prices from world market prices, the effects of export restrictions on both domestic prices and world market prices have to be taken into account. Martin and Anderson (2012) estimate that almost 30%[6] of the increase in world wheat prices from 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.
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), i.e., that governments were not successful in stabilising domestic prices. From a global point of view, the domestic wheat price in Ukraine would have increased similarly if

[6] For comparison, Anderson and Nelgen (2012a) estimate that 19% of the world wheat price increase was caused by price insulating behaviour.

no country worldwide had engaged in price insulating behaviour from 2006-2008 – 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 incurred by loaded ships locked in harbours, and the losses resulting from delayed or reduced investments in the Ukrainian grain production sector.

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

7 Literature

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