The Law of One Price under State-Dependent Policy Intervention:

An Application to the Ukrainian Wheat Market

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Abstract

This paper investigates the effects of state-dependent policy interventions on price transmission. Our empirical application focuses on price linkages between the Ukrainian wheat price and the world price. The empirical analysis is based on the smooth transition conintegrating (STC) framework and follows the general procedures used to investigate long-run equilibrium and short-run error correction. The results indicate that there is regime-switching behavior in the long-run relationship between the Ukrainian and world markets, conditional on the world price. When the world price of wheat is below the threshold of $185/ton, the transmission elasticity of domestic price with respect to the world price approaches unity. However when the world price is above the threshold level, the transmission elasticity drops to 0.7. Finally, we also find that adjustments toward the long-run equilibrium take place through changes in Ukrainian domestic price alone. Our results suggest that the Ukrainian wheat market is well integrated into the world market. However, government intervention can cause significant long-term losses for Ukrainian producers.

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1 Introduction

The law of one price (LOP) is one of the fundamental principles of trade theory. It states that homogeneous goods sold in different regions will sell at the same price when expressed in the same currency. The LOP has been considered as an important indicator of market efficiency because it illustrates to what extent markets are linked across space. When referring to economies in transition, the LOP is also an important index of the market liberalization.

A rich body of the empirical economics literature has investigated the LOP among spatially separated markets. Early studies use correlation coefficients and regression techniques to directly test the equality of prices in different regions (e.g., Isard 1977, Richardson 1978, Protopapadakis and Stoll 1986). The results usually do not support the LOP. Some economists blamed transaction costs (primarily transportation costs) for the failures of prices to converge. They thus proposed a modified version of the LOP which stated that prices of homogeneous products in any two locations should not differ by more than the costs of transferring those goods from one location to the other. As long as trade can take place freely, price shocks in one region can be buffered and price co-movement between regions will be observed. Several empirical studies (e.g., Goodwin 1992 and Michael et al. 1994) have incorporated transaction costs into the analysis and found some supportive evidence.

Modern empirical studies have noticed the nonstationary attribute of the price data and proposed a different framework for testing the LOP. Engle and Granger (1987) point out that, given a pair of first-order integrated series, if there is a linear combination between them which is stationary, the two processes are said to have a long-run equilibrium or simply are said to be cointegrated. Their approach has provided researchers of price transmission (spatial and vertical) with valuable tools for jointly modeling and drawing inferences about the long-run price relationship, together with the short-run adjustments toward the equilibrium.

Some economists (e.g., Goodwin and Piggott 2001) suggest that given cointegration, the short-run adjustments to the equilibrium may not be linear because of the transaction costs associated with arbitrage. Deviations from long-run equilibrium within the transaction cost band will not trigger any adjustment simply because it is not profitable to do so; but deviations that fall outside of the band will trigger trade activities and thus should be mean reverting. This validates the introduction of nonlinear regime-switching autoregressive models and the corresponding (vector) error correction (EC) models into the analysis. Following this idea, an extensive literature has investigated price transmission accounting for nonlinear adjustments by using various versions of regime-switching EC models (i.e., threshold EC, smooth transition EC, and Markov-switching EC). Under this framework, supportive evidence for the LOP have been reported by Lo and Zivot (2001), Sephton (2003), Balcombe, Bailey, and Brooks (2007), Park, Mjelde, and Bessler (2007), and Goodwin, Holt, and Prestemon (2011).

An assumption underlying this transaction-cost version of the LOP, and therefore the use of error correction models, is that trade is free and open (i.e., without barriers, such as tariffs, quotas, or regional arbitrage interventions). However, trade restrictions do often exist, especially when dealing with agri-food markets. Policy interventions may not only affect short-run dynamic adjustments, but may also alter or even eliminate any long-run market integration under certain conditions. Export/import taxes, subsidies, quotas, certificates, and direct bans, can create a considerable wedge between world and domestic prices, and thus lead to incomplete transmission in prices.

As a measure of transaction costs, direct quantification of policy intervention is difficult. Policy intervention often reflects a state-dependent reaction rather than a constant behavior. For instance, if the objective of policy active exporting country is to stabilize the domestic price, export controls might be triggered when the world price is too high, and subsidies would be applied when the world price is too low. This state-dependent feature indicates a nonlinear relationship between prices. Although the extension of the concept of cointegrating relationship to a nonlinear framework is not new (see Park and Phillips 1999, 2001, Chang and Park 2003, Saikkonen and Choi 2004, Gonzalo and Pitarakis 2006, among others), the procedure to test and estimate nonlinearity in cointegrating vectors is. The policy effects therefore are often investigated indirectly by, on one hand, adding dummies or conducting investigations in different time periods (e.g., Thompson, Sul, and Bohl 2002 and Baffes and Ajwad 2001) and adding a constant term (sometimes together with a proportional term) in the price transmission equations, to account for a fixed policy effect (e.g., Mundlak and Larson 1992).

The objective of this paper is to provide an investigation of the effects of state-dependent policy intervention on spatial price transmission. In pursuing this objective, this study contributes to the literature in three ways. First, we relax the linear cointegrating restriction and allow the long-run equilibrium to be nonlinear based on the state of intervention. Second, we also allow the short-run error correction processes to differ by state, conditional on nonlinearity in the long-run price relationship. Third, we propose an empirical application related to the Ukrainian wheat market. We investigate the price linkages between the Ukrainian and the world markets. Ukraine is an interesting case study, as it is a typical transition country with active and frequent government intervention. It is also one of the world's top grain exporters. Appropriate investigations of integration between the market and the world market (if any) will provide valuable information for future policy recommendations regarding food security, market efficiency, and trade liberalization.

The remainder of the paper is structured as follows. Section 2 outlines the conceptual framework which introduces the state-dependent policy intervention into the price transmission analysis and develop a simple regime-switching LOP framework. Section 3 provides a brief background on the Ukrainian wheat market and relevant trade policies used over the sample period. Section 4 is dedicated to the empirical procedure, followed by a presentation of the results. Section 5 discusses the policy implication and Section 6 concludes the study.

2 Conceptual Framework

The model below builds upon earlier efforts of Mundlak and Larson (1992). We expand their work by introducing the state-dependent feature of policy intervention into the model. It thus allows price linkages to exhibit regime-switching behavior. To empirically investigate the relationship between the domestic and world prices in the presence of policy intervention, Mundlak and Larson (1992) propose the following model

$$P_{it} = P^*_{it} E_t S_{it} \tag{1}$$

where P_{it} denotes the domestic price of commodity i at time t. According to the LOP, it can be expressed as a product of the world price P^*_{it} , the nominal exchange rate E_t, and the policy intervention S_{it}. This study does not investigate exchange rate transmission issues and focuses on the linkages between the two prices that are measured in the same currency (US$ in our case), which is a common feature of internationally traded commodities. When rewriting the price relation equation in the logarithmic form, we obtain

$$p_{it} = p^*_{it} + s_{it} \tag{2}$$

where $p^*_{it} = \ln(P^*_{it} E_t)$. Assuming policy depends on world market conditions, Mundlak and Larson (1992) propose the following policy reaction relationship

$$s_{it} = \phi_i + \pi_i p^*_{it} \tag{3}$$

where π is a policy reaction index which reflects to what extent the government reacts to world market price. Combining (2) and (3), for a given homogenous commodity i, the domestic and world price relationship can be expressed in logarithmic form

$$p_{it} = \phi_i + (1 + \pi_i) p^*_{it} \tag{4}$$

We expand Mundlak and Larson's (1992) work by letting the policy intervention equation in (3) be a state-dependent reaction function which itself is induced by world market conditions

$$s_{it} = \begin{cases} 0 & \text{if } \theta_1 < p^*_{it} < \theta_2, \\ \phi_1 + \pi_{i1} p^*_{it} & \text{if } p^*_{it} \le \theta_1, \\ \phi_2 + \pi_{i2} p^*_{it} & \text{if } p^*_{it} \ge \theta_2. \end{cases} \tag{5}$$

Substituting (5) into (4), we then obtain the corresponding state-dependent price linkage as

$$p_{it} = \begin{cases} k_i + p_{it}^* & \text{if } \theta_1 < p_{it}^* < \theta_2, \\ \phi_{i1} + (1+\pi_{i1})p_{it}^* & \text{if } p_{it}^* \leq \theta_1, \\ \phi_{i2} + (1+\pi_{i2})p_{it}^* & \text{if } p_{it}^* \geq \theta_2. \end{cases} \tag{6}$$

In logarithmic form, the econometric specification can be written as

$$p_{it} = (c_1 + \rho_1 p_{it}^*)I_1 + (c_2 + \rho_2 p_{it}^*)I_2 + (c_3 + \rho_3 p_{it}^*)I_3 + \varepsilon_{it} \tag{7}$$

where c_i, $i = 1, 2, 3$ are the constant terms that can be interpreted as an overall effect of a set of factors affecting price signals, including transportation costs, the degree of product homogeneity, changes of the consumer or producer price indexes, and the fixed part of policy effects as shown in (5), and so on. The term ε_{it} is a stationary disturbance and I_i, $i = 1, 2, 3$ are indicator functions which satisfy the conditions that the world price is within a certain range, or below or beyond a certain threshold. Again, assume the purpose of government intervention is to stabilize the domestic price, as long as the world price is staying within a certain range, let's say, a commodity-specified "open trade band", the government will not (actively) intervene in trade and the open trade assumption holds. However, if the world market price goes outside the band, either by becoming too low or too high, the government will intervene. Under these circumstances, as long as the world price is still within the band, any fluctuations of the world price would not trigger government intervention, thus one can expect a close-to-unity price transmission elasticity from the (conintegrating) regression in (7). However, if the world price goes below the lower threshold, export subsidies might be introduced to maintain a relatively high and stable domestic price and to support the domestic producers. In this case, a positive π_1 is expected, thus a greater-than-unity, price transmission elasticity (i.e., $\rho_2 > 1$) is also expected, if the LOP holds. Conversely, when the world price is "too high" and beyond the upper threshold value θ_2 , the government intervenes through the introduction of export taxes, bans, and/or quotas to lower the domestic price. A less-than-unity coefficient ρ_3 would be expected.[1]

[1]It is worthwhile to mention that, in reality, direct government interventions to domestic markets/prices may not occur in developed counties, but are not rare in those less developed counties and economies in

3 Ukrainian Wheat Market and World Food Crisis

Ukraine is the second largest European country after Russia. It became independent when the Soviet Union dissolved in 1991. The economy experienced a large increase in GDP growth after an eight-year recession that immediately followed the dissolution. It is a globally important grain supplier largely due to its endowment arable land. Ukraine has more than 100 million acres of cropland and permanent pasture with fertile soils–approximately 40% of the world's black soils, year-round ice-free ports, and proximity to key import markets in the Middle East, Northern Africa, and the European Union (von Cramon-Taubadel and Zorya 2001). Though grain production suffered from dramatic declines in the first decade following independence, output has considerably increased since then. In marketing year (MY) 2009/10, Ukraine was easily among the world's top three leading grain exporters (after Brazil and Russia). Between 2008 and 2010, Ukraine, together with Russia, exported an average of 29 million tons of wheat annually. This accounted for 21.3% of world wheat exports and was greater than the exports of any of the other major exporters US, Canada, EU-27, and Australia (Goychuk and Meyers 2011).

Although Ukraine is a large grain exporter, it is still plagued by food security issues. As pointed out by von Cramon-Taubadel and Zorva (2001), food consumption of an individual (or a country) does not just depend on its production ability, but more importantly, on his/her endowments, working capacity, and exchange entitlements (i.e., the ability to exchange these endowments for food). Even if a country is a net exporter of food, its vulnerable, low-income groups can still suffer from hunger. In Soviet times, the economy of Ukraine was the second largest in the Union and was an important industrial and agricultural component of the country's planned economy. With the dissolution of the Soviet system, the country moved from a planned economy to a market economy. The transition was difficult, and plunged the majority of the Ukrainian people into poverty. A large part of the population could not afford food, and some had to rely on a subsistence diet of bread

transition. Some reasons are: lack of trade and economics knowledge, traditions of mixed and/or planned economy, poor infrastructure system, tight budgets, and less developed social welfare supportive programs.

and tea (von Cramon-Taubadel and Zorya 2001). As a result, rising food prices are most likely to incite political unrest and violence. Given the political sensitivity of food prices, combined with Ukraine's history of a planned economy, the Ukrainian government always reacts quickly to the global rise in grain prices. In Ukraine, grain markets are often considered as a "political tool". Both local and central governments control crop and food prices (Brmmer, von Cramon-Taubadel, and Zorya 2009). When world wheat prices soar, the response of Ukrainian government is often populist in nature. The government often accuses traders/speculators driving up wheat prices. As a result, they introduce export certification, export quotas, and fixed bread prices to try to control the market prices.

World food prices increased dramatically in MY 2007/2008, creating a global food crisis. In 2008, U.S. wheat export prices rose from \$375/ton in January to \$440/ton in March, and Thai rice export prices increased from \$365/ton to \$562/ton. This came on the heels of a 181% increase in global wheat prices over the 36 months preceding February 2008, and an 83% increase in overall global food prices over the same period (Revenga 2011). Similarly, since July 2010, prices of many crops have risen significantly. World food prices reached a historic peak in January 2011, exceeding prices reached during the food crisis of MY 2007/08. Corn increased by 74%; wheat prices went up by 84%; and sugar prices by 77% (Oxfam online 2011).

Food price crisis caused political and economic instability in Ukraine. In both periods, the initial response of the Ukrainian government to rising food prices was to implement grain export controls, primarily by issuing export quotas. The argument behind these market interventions is that they are needed to guarantee food security and protect domestic consumers from rising international food prices. The first export quotas were introduced in late September 2006. The quota volumes set for the MY 2007/08 were especially low. They virtually banned exports over a certain time period.[2] In July 2008, export quotas were cancelled due to the gradual decreases of world market prices and a large domestic grain harvest in MY 2008/2009. In addition, Ukraine had an obligation to cancel the export restrictions as

[2]The total export quota in MY2007/08 is 1.2 million tons, compare to a 12.9 million tons net export in MY2008/2009.

part of its WTO commitments.[3] In October, 2010, the Ukrainian government again enacted a resolution requiring quotas and licenses for exporting grain. While the protectionist policy came under attack from both foreign and domestic observers, the government extended the export grain quotas until June 30, 2011. Moreover, the government used corrupt practices of allocating export quotas and licenses wherein an unknown company Khlib Investbud received the majority share and gained market power in the grain export industry. In place of the quota, contract price export duties of 9% of the contract price were introduced on July 1, 2011 and remained in effect until January 1, 2012. At the same time, although direct government intervention in the grain markets is common in Ukraine when the market price is "too high," the government does not subsidize grain exports when the world market price is low.

Following the above description, we thus propose to use a two-regime policy response model – "inactive" intervention and "active" intervention–based on the world market price– to evaluate the relationship between domestic and world wheat prices. In particular, when the world market price falls below a certain threshold, no significant export controls are triggered. We thus expect a near-unitary price transmission elasticity of the domestic price respect to the world market price, if the LOP holds. However, if the world market price reaches and exceeds the threshold, export controls would be triggered. Accordingly, increases of the world market price would not fully pass-along to the Ukrainian domestic price and a less-than-unity transmission elasticity can be expected.

Based on this information, a two-regime threshold cointegrating regression model is appropriate to model the price linkages. However, the threshold models are based upon the assumption that the transition from one regime to another is abrupt and discontinuous. If threshold models are used to capture the policy-switching behavior, the break between regimes can only be sharp and discontinuous if any policies can be fully carried out instantly without any delay. However, both policy intervention and market adjustment take time and would probably develop gradually for a while before any changes can be made. Therefore

[3]Ukraine became the WTO's 152nd member on May 16, 2008.

the regime-switching behavior of the price transmission is likely to be smooth. A smooth transition conintegrating regression model is thus utilized in the empirical stage. It is important to note that such a model specification also allows for rapid adjustment, such as that imposed by discrete threshold models.

4 Data and Empirical Procedure

This study uses weekly observations for the world market and Ukrainian 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 FOB price of wheat (classification other wheats) in Rouen, France (obtained from consulting company HGCA 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 1 shows the Ukrainian domestic and world wheat price series. Figure 2 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 export/import controls 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 behaviors for the U.S. and German situations. This suggests an impact on price linkages resulting from government intervention.

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).[4] Table 1 presents the test results. The ADF tests fail to

[4]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 the ADF results.

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.

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 labor 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 a threshold behavior 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 commodity 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); [5]
3. Test stationarity using the residuals obtained from the estimated STC model;

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

4. Test linearity versus nonlinearity for error correction procedures, again using 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.

4.1 Test Linear Versus STC Long-Run Relationship

Consider a smooth transition cointegrating (STC) model

$$y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + z_t, \quad t = 1, 2, \ldots, T \tag{8}$$

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 function

$$g(x_t - c; \gamma) = \left[1 + \exp(-\gamma(x_t - c))\right]^{-1} \tag{9}$$

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 (8) 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 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 (8). That is, $H_0 : \alpha_2 = \beta_2 = 0$, whereas under the alternative hypothesis of H_1: at least one of α_2 and $\beta_2 \neq 0$. 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 (9). 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 . 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 (8) is small. We thus adopt the third-order Taylor approximation and rewrite the transition function as

$$g(x_t - c; \gamma) \approx b\gamma(x_t - c) + d[\gamma(x_t - c)]^2 + h[\gamma(x_t - c)]^3 \tag{10}$$

The testing procedure involves estimating the corresponding auxiliary regression using OLS[6]

$$\begin{aligned} y_t &= \alpha_1 + \alpha_2 \Big\{ b\gamma(x_t - c) + d\big[\gamma(x_t - c)\big]^2 + h\big[\gamma(x_t - c)\big]^3 \Big\} \\ &\quad + \beta_1 x_t + \beta_2 x_t b\gamma(x_t - c) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} \\ &= \omega + \phi_1 x_t + \phi_2 x_t^2 + \phi_3 x_t^3 + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \quad t = K+1, \ldots, T-K \end{aligned} \tag{11}$$

The null hypothesis of linearity is $\phi_2 = \phi_3 = 0$. The LM statistic is $\tau = \hat{\Phi}'[\hat{\sigma}_\varepsilon^2 (M^{-1})_{xx}]^{-1}\hat{\Phi}$, where $\hat{\Phi} = [\hat{\phi}_2 \;\; \hat{\phi}_3]'$ are the OLS estimates of $[\phi_2 \, \phi_3]$, $\hat{\sigma}_\varepsilon^2$ 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,

[6]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 a third-order approximation only for the transition of the intercept term and using the first-order approximation for the transition involving the regressors.

$\tau \xrightarrow{d} \chi^2(p+1)$, where p (1 in our case) is the dimension of the model. Test results are presented in Table 2.

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 Germany 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 prior expectation since these two countries have not implement 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.

4.2 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, Tersvirta, and Franses 2002 and Enders 2010) usually suggest using a nonlinear least square (NLLS) technique to obtain the estimates of the parameters in (8).[7] The estimate of the parameter vector $\theta = [\gamma\, c\, \alpha_1\, \alpha_2\, \beta_1\, \beta_2]$ will satisfy

$$\hat{\theta} = \arg\min_{\theta} Q_T(\theta) = \sum_{t=1}^{T} \left[\tilde{y}_t - y_t(x_t;\theta)\right]^2 \tag{12}$$

where $\tilde{y}_t$ is sample observations and $y_t(x_t;\theta)$ is the so-called skeleton of the model given in (8). 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 (12) is consistent, the asymptotic distribution involves a bias if regressors and error are dependent, which makes the above NLLS estimator inefficient and unsuitable for use in

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

hypothesis testing. They thus propose a GaussNewton (G-N) type estimator that utilizes the NLLS estimator obtained from (12) 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 (12) and then to use $\hat{\theta}$ as the initial value and estimate the following augmented STC model

$$y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \quad t = K+1, \dots, T-K \tag{13}$$

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 $\theta_1, \theta_2, \dots$ 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 that 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, Tersvirta, 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 9 with a normalized version

$$g(x_t - c; \gamma) = \Big[1 + \exp(-\frac{\gamma}{\hat{\sigma}_x^2}(x_t - c))\Big]^{-1} \tag{14}$$

Table 3 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 4. 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.[8]

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 regressor-error 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.

$$\begin{aligned} \hat{y}_t &= \begin{cases} -0.86 + 1.14x_t & \text{if } g = 0, \\ 1.27 + 0.70x_t & \text{if } g = 1, \end{cases} \\ \text{and} \quad g(x_t - c; \gamma) &= 1/\big\{1 + \exp\big[-3.87(x_t - 5.21)/0.16\big]\big\}. \end{aligned} \tag{15}$$

The results confirms a regime-switching behavior in the long-run relationship between Ukrainian and world prices, based on the level of world market prices. The estimated threshold value for the transition variable is 5.2 in logarithms, or \$185. When the world

[8]Following the suggestion of David Dickey, we also test the unit root and stationarity of the residuals by regime. In particular, we split the residuals into two group 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.

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 3. 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.

4.3 Short-run Dynamic Adjustment

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

$$\Delta z_t = \begin{cases} \phi_1 z_{t-1} + \varepsilon_1 & \text{if } z_{t-1} < \theta_1, \\ \phi_2 z_{t-1} + \varepsilon_2 & \text{if } \theta_1 < z_{t-1} < \theta_2, \\ \phi_3 z_{t-1} + \varepsilon_3 & \text{if } z_{t-1} > \theta_2. \end{cases} \tag{16}$$

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

$$\Delta y_t = \begin{cases} \sum\limits_{i=1} \alpha_{1i}\Delta y_{t-i} + \sum\limits_{j=1} \beta_{1j}\Delta x_{t-j} + \varphi_1 z_{t-1} + e_1 & \text{if } z_{t-1} < \theta_1, \\ \sum\limits_{i=1} \alpha_{2i}\Delta y_{t-i} + \sum\limits_{j=1} \beta_{2j}\Delta x_{t-j} + \varphi_2 z_{t-1} + e_2 & \text{if } \theta_1 < z_{t-1} < \theta_2, \\ \sum\limits_{i=1} \alpha_{3i}\Delta y_{t-i} + \sum\limits_{j=1} \beta_{3j}\Delta x_{t-j} + \varphi_3 z_{t-1} + e_3 & \text{if } z_{t-1} > \theta_2. \end{cases} \tag{17}$$

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.[9]

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

$$z_t = \alpha_1 z_{t-1} + e_t, \text{ and} \tag{18}$$

$$z_t = \alpha_1 z_{t-1} I_1(\gamma) + \alpha_2 z_{t-1} I_2(\gamma) + e_t, \text{ respectively.} \tag{19}$$

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, $\hat{\gamma} = \arg\min SSR_2(\hat{\gamma})$. The search over all possible values of the

[9] 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 aside, Ukraine is a pure wheat exporter in our sample time period.

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

$$F_{12} = n(SSR_1 - SSR_2^{\min})/SSR_2^{\min} \tag{20}$$

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 $\Delta_k y_t = y_t - y_{t-k}$, $k = 1, 2, \ldots, k_{\max}$ 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,

$$z_t = \alpha_1 z_{t-k} + e_t \text{ vesus } z_t = \alpha_1 z_{t-k} I_1(\gamma) + \alpha_2 z_{t-k} I_2(\gamma) + e_t. \tag{21}$$

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 5. Model selection is based on AIC and SBC.

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 6. 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 6.[10] 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.

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

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

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 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.

5 Policy Implications

This paper uses a more flexible STC model to investigate the price relationship between Ukraine and world markets, taking the state-dependent trade intervention into account. We find that a long-run equilibrium relationship exists and varies according to the world price. When the world price is below a certain threshold, Ukrainian and world markets are well integrated. However, when the world price exceeds the threshold level, it triggers active interventions, and the two markets are less integrated. In particular, only 70% of changes of the world price would be transmitted to Ukrainian price. In other words, 30% of potential export revenues are lost, other things being equal. The regime-switching long-run equilibrium provides a framework to estimate and predict the potential domestic export loss under certain scenarios. For example, consider the average world price during January 2010 to September 2011, \$213.8/ton. Assume further that the reduced export quantity is 10 million tons. Then, a 50% increase in world price will result in a \$320.7 ($213.8 \times 10 \times 0.5 \times 0.3$) million revenue loss for the domestic growers. What makes the domestic producers lose even more is that on the input side, rising energy prices in recent years have influenced the costs of production and trade. Production revenues have been further reduced for Ukrainian producers as their production depends on importing energy from Russia and fertilizers from international markets.

In summary, the two-regime long-run price transmission results indicate that the Ukrainian market itself is well integrated with the world market. However, continuous government interventions in trade activities can cause significant losses for the domestic producers in the long run.

To give a more complete story of the impacts of export controls on Ukrainian domestic economy, we briefly discuss some important findings from other studies, in the hope of offering some suggestions for future policy recommendations. According to recent studies (e.g., von-Cramon and Raiser 2006 and Brmmer, von Cramon-Taubadel, and Zorya 2009), although the stated purpose of these export controls is to help those low income consumers, these are the people who actually benefit the least from the quota. First, wheat prices contribute only a certain percentage to the final bread price. The impact of lower wheat prices on the prices of meat and dairy is quite limited. Second, though wheat prices have been somewhat controlled, prices for flour and bread have actually risen since the introduction of the quote in 2006. Instead of the poor consumers, flour millers and animal feed producers, whose profit margins increase as a result of falling grain prices on the domestic market, are the main beneficiaries of the quota.

The quota system has also imposed big losses on international agribusiness companies and traders that have invested billions of dollars in farming, trading, storage, processing and export facilities. Furthermore, some have argued that the government used corrupt practices of allocating export quotas and licenses which resulted in unfair and nontransparent competitions in the trade market which hurt the majority of traders.

The future policy implications, in this paper are in accordance with von-Cramon and Raiser report (2006) which argues:

> "The quota system is ineffective (does not reach the poor), inefficient (imposes large cost for very limited gain), and led to corruption. The suggestion is therefore to abolish the quota system as soon as possible .... Alternative measures including the use of means tested cash transfers need to be considered to protect the poor from rising food prices."

6 Conclusion

The extent and magnitude of policy intervention on price transmission, when allowing for state-dependent attributes, on price transmission offer valuable information on price linkages and market integration. More generally, state-dependent or regime-switching long-run price equilibrium can result from other factors, such as state-dependent exchange rate pass-through, market power, and/or asymmetric information. It is thus a useful extension and generalization of linear cointegration approaches for modeling price transmission that has appeared in the literature. However, the development of nonlinear cointegration techniques and their application to price transmission are both novel and deserve more attention.

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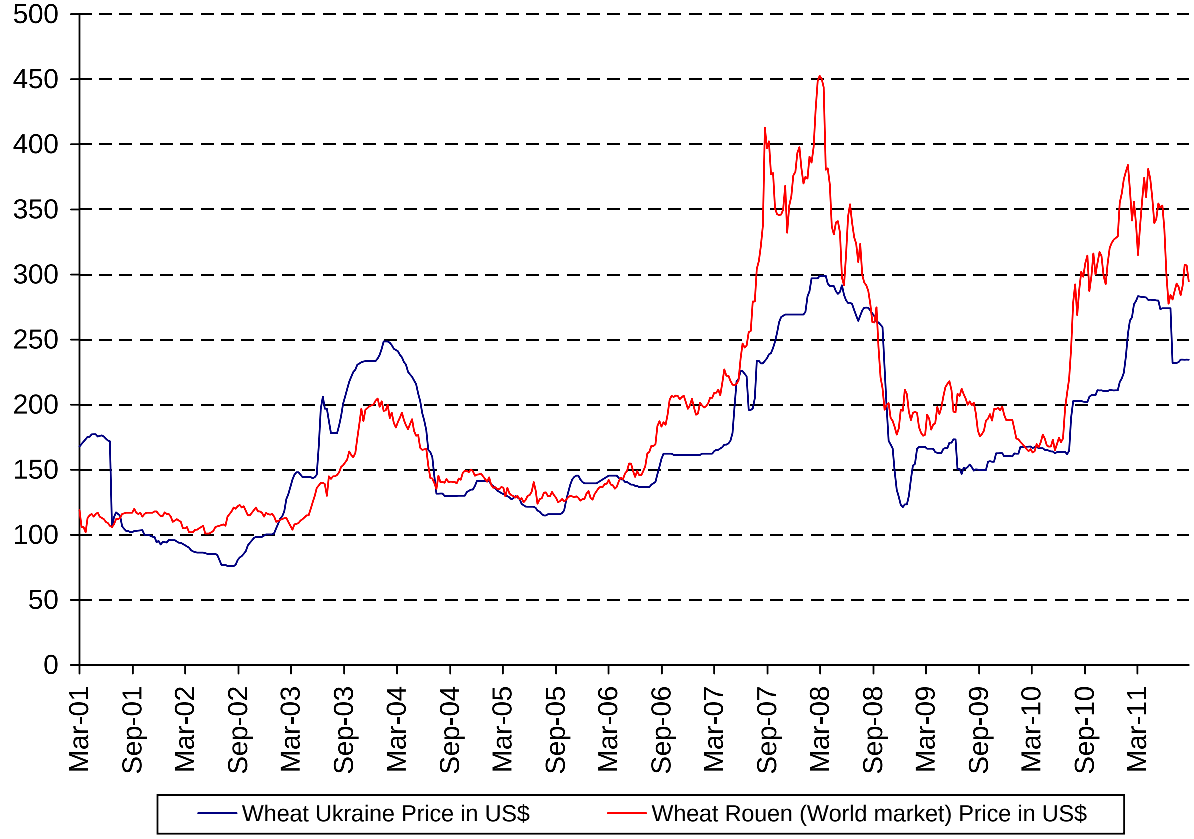
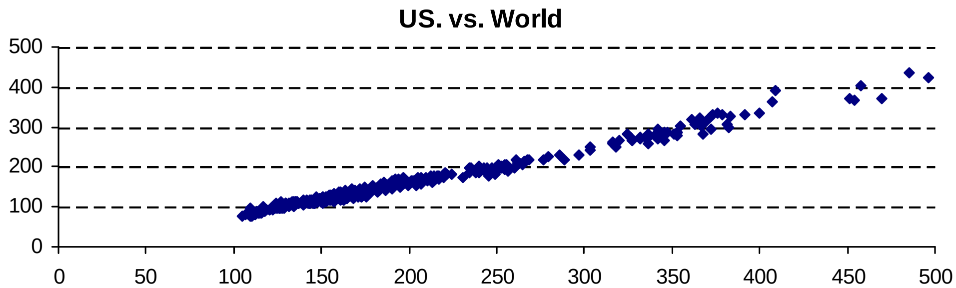
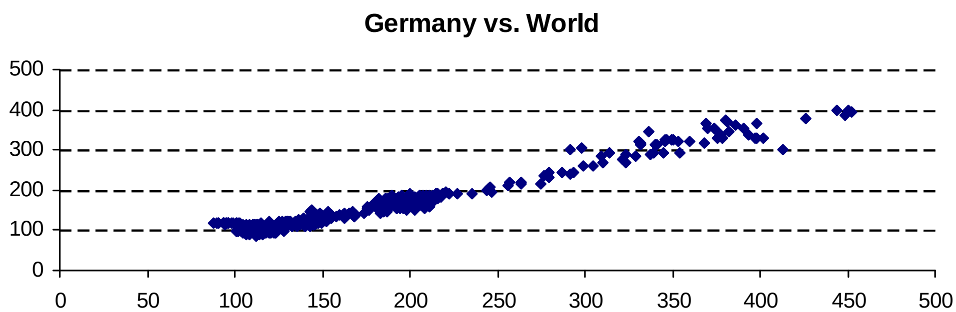


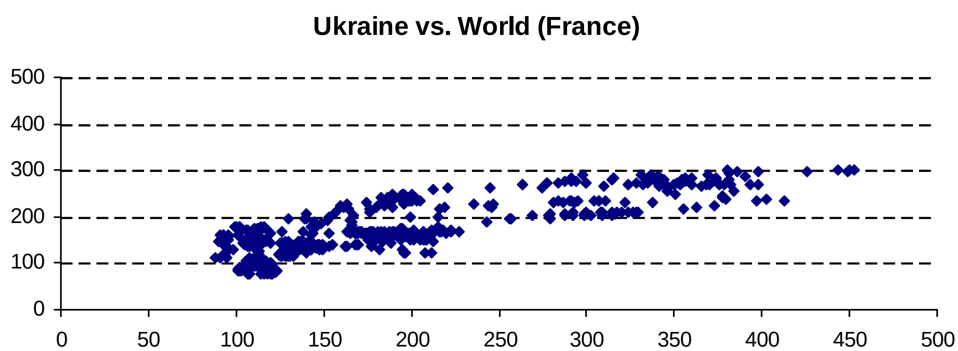
Figure 1: Ukrainian and World Market Prices (US$/ton): March 16, 2001-September 9, 2011



(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 2: Domestic and Its Corresponding World Market Prices.

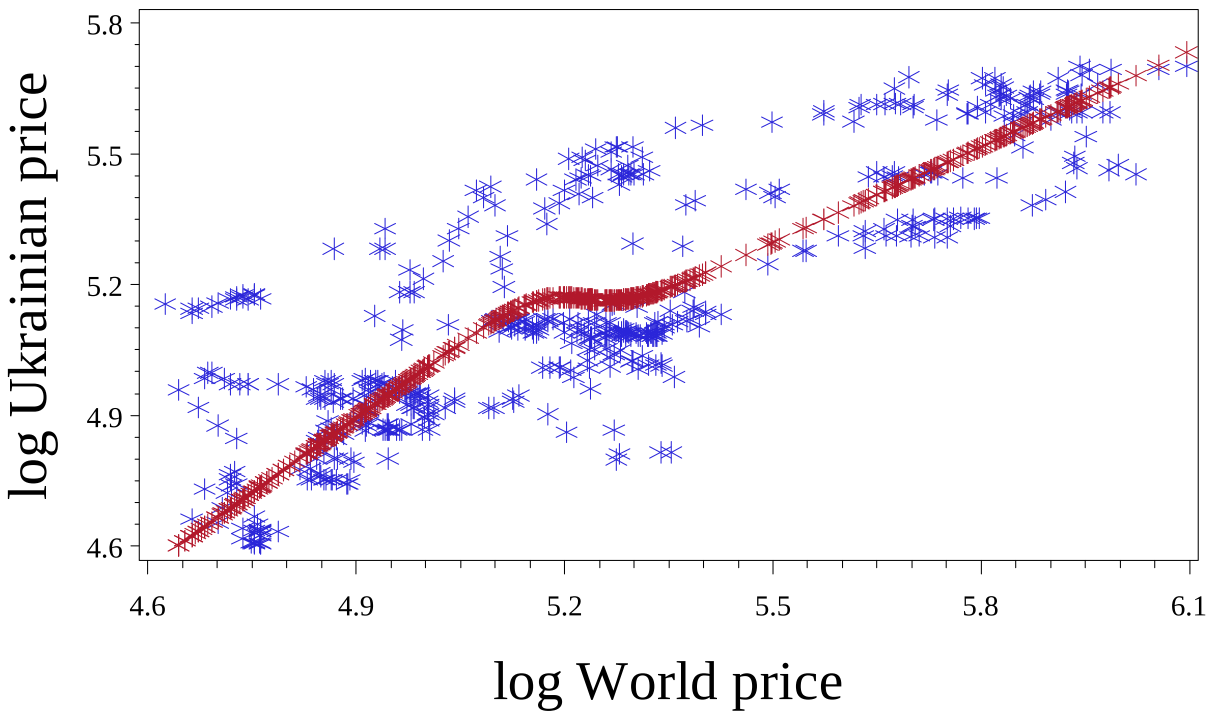


Figure 3: Smooth Transition Cointegrating Model Fit

Table 1: Unit Root Tests for Price Data (in natural logarithms)

	Lags	World price	Lags	Ukraine price	Lags	ΔWorld price	Lags	ΔUkraine price
Dickey-Fuller								
Single Mean	Lags		Lags		Lags		Lags	
ρ	3	-3.15	3	-6.81	3	-319.74	3	-252.802
$\Pr < \rho$		0.64		0.29		(< 0.001)		(< 0.001)
τ_μ	3	-1.19	3	-1.75	3	-9.95	3	-9.23
$\Pr < \tau_\mu$		0.68		0.4		(< 0.001)		(< 0.001)
Trend								
ρ	6	-8.54	3	-12.98	6	-319.78	6	-8.54
$\Pr < \rho$		0.54		0.26		(< 0.001)		(< 0.001)
τ_μ	6	-2.05	3	-2.66	6	-9.94	6	-253.78
$\Pr < \tau_\mu$		0.58		0.25		(< 0.001)		(< 0.001)
KPSS								
Single Mean	6	4.81	6	2.93	6	0.07	6	0.1
Trend	6	0.3	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 2: Linear vs. Smooth Transition Cointegrating Vector Tests

	Ukrainian vs. world market price	United States vs. world market price	German vs. world market price
Lags and Leads		Statistic τ	
$\sum_{j=-K}^{K} \alpha_j \Delta p_j^{wd}$		(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

Note: 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 3: Estimates of the Smooth Transition Cointegrating Models.

	STC, no lags and leads			STC, with lags and leads		
Parameter	Estimate	Approx Std Err	Approx $Pr > \lvert t \rvert$	Estimate	Approx Std Err	Approx $Pr > \lvert t \rvert$
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	0.5	< 0.01
α_2	2.13	0.67	< 0.01	2.77	0.69	< 0.01
β_1	1.14	0.1	< 0.01	1.19	0.1	< 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 4: 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 5: 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
$k = 2$	0.9	0.45	0.42
$k = 3$	0.9	0.44	0.44
$k = 4$	0.92	0.46	0.42

Table 6: Estimates for Linear ECMs

	All residuals		Residuals from		Residuals from	
	All residuals		regime1		regime2	
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	0.1	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
Δx_{t-1}						
Δx_{t-2}	-0.14	0.053				
Δx_{t-3}	-0.13	0.054	-0.2	0.095		
Half-life	17.7wks		17.7wks		8.0wks	
AIC	-294.92		-207.13		-193.5	
SBC	-252.42		-174.2		-182.99	
Observation	542		250		292	