E C O N O M I C S B U L L E T I N

Asymmetric price adjustment of Ukrainian feed wheat export prices in relation to U.S. maize exports: A Note

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Abstract

This note aims to investigate the price dynamics between Ukrainian feed wheat exports and U.S. maize exports employing the method by Enders and Siklos (2001) to allow for threshold adjustment. The analysis reveals slow amounts of adjustment towards equilibrium for a positive change in the equilibrium relationship, but a substantial amount of adjustment for a negative change in the equilibrium relationship. This would imply that Ukrainian prices would have to decrease at a faster rate in order to retain their market share. The fact that there is some evidence of asymmetric price adjustment is not surprising given that Ukraine is a minor exporter in relation to the U.S. and that policy intervention exists in the Ukrainian wheat market. The nature of price dynamics sheds light on how Ukraine has been strategically responding to U.S. maize exports since the mid 1990s.

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1. Introduction

During the era of the Soviet Union, most of the wheat produced by Ukraine used to go to Russia and other FSU countries. Since Ukraine gained independence in 1991, policy reforms aimed towards a market based economy have been slow. Though prices were liberalized after 1992, agricultural subsidies were in place contributing to budget deficits and inflation (Agriculture and Agri-Food Canada 2003). Since the mid-1990's, increasing amounts of wheat are being exported to other countries across the world. The wheat market is highly politicized and is considered to be strategic in Ukraine with policy makers actively intervening on agricultural markets (Brummer *et. al.,* 2006). In the case of feed wheat exports, Ukraine competes with the U.S. maize exports for markets in South Korea and Israel. The low production costs of wheat in Ukraine and the subsequent currency devaluation that took place in 1999 has enabled Ukrainian traders to offer lower prices and increasingly undercut the prices set by the U.S. (USDA 1999).

The prices charged by Ukrainian exporters of feed wheat are expected to move closely with U.S. maize export prices over time due to arbitrage or substitution or both. In other words, the export prices of these closely related commodities will not diverge too far from each other if a price change by one major exporter is to be followed by a gradual similar price change by the other major exporter. In this way, by examining the relationships between prices over time, the relevant market can be defined. From an econometric point of view, this would imply that prices of the commodity should be cointegrated. The presence of government intervention allows the possibility of developing strategic responses to price adjustment. A possible strategy might be to match a competing country's price decreases but not increases. Asymmetric adjustment may occur when the response of competing prices to a price fall is not equal to that of a price rise (Ghoshray 2007). Recent developments in cointegration techniques have allowed for cointegration with threshold and momentum-threshold adjustment.

This note aims to investigate the price dynamics between Ukrainian feed wheat exports and U.S. maize exports employing the method by Enders and Siklos (2001) to allow for asymmetric adjustment.[1] Given that Ukraine has been making a transition to a market based economy since gaining independence in 1991, this study on how Ukraine feed exports of wheat compete with U.S. maize exports is both significant and timely. The next section describes the econometric model, followed by a discussion of the results. Finally, the last section concludes.

2. Econometric Model

At the onset, the Engle and Granger (1987) method is employed to test for cointegration. The long-run relation of the two export prices given by the equation below is estimated using ordinary least squares:

$$P_t^{UKR} = \alpha + \beta P_t^{US} + \varepsilon_t \qquad (1)$$

where P_t^{UKR} and P_t^{US} are non-stationary I(1) prices of Ukraine and the U.S. respectively. The arbitrary constant α accounts for transfer costs, β denotes the price transmission elasticity and ε_t is the error term which may be serially correlated. The following Dickey-Fuller test on the estimated residuals $\hat{\varepsilon}_t$ of (1) allows us to determine whether the prices adjust to any deviation from the long run equilibrium by testing the null hypothesis $(H_0 : \gamma = 0)$ of no cointegration in the equation below:

$$\Delta\hat{\varepsilon}_t = \gamma\hat{\varepsilon}_{t-1} + \omega_t \tag{2}$$

where ω_t is a white noise error term.[2]

However, Enders and Siklos (2001) argue that the test for cointegration and its extensions are mis-specified if adjustment is asymmetric. They consider an alternative specification, called the threshold autoregressive (TAR) model, such that (2) can be written as:

$$\Delta\hat{\varepsilon}_t = I_t\gamma_1\hat{\varepsilon}_{t-1} + (1-I_t)\gamma_2\hat{\varepsilon}_{t-1} + \omega_t \tag{3}$$

where I_t is the Heaviside indicator function such that:

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{\text{t-1}} \geq \tau \\ 0 & \text{if } \hat{\varepsilon}_{\text{t-1}} < \tau \end{cases} \tag{4}$$

This specification allows for asymmetric adjustment. If the system is convergent, then the long run equilibrium value of the sequence is given by $\hat{\varepsilon}_t = \tau$, where τ is the estimated threshold. A method of searching for a consistent estimate of the threshold was undertaken by using a method proposed by Chan (1993).[3] The sufficient conditions for the stationarity of $\hat{\varepsilon}_t$ are $\gamma_1 < 0$, $\gamma_2 < 0$ and $(1+\gamma_1)(1+\gamma_2) < 1$ (Petrucelli and Woolford 1984). In this case if $\hat{\varepsilon}_{t-1}$ is above its long run equilibrium value, then adjustment is at the rate γ_1 and if $\hat{\varepsilon}_{t-1}$ is below long run equilibrium then adjustment is at the rate γ_2. The adjustment would be symmetric if $\gamma_1 = \gamma_2$. However, if the null hypothesis $H_0 : (\gamma_1 = \gamma_2)$ is rejected then using the TAR model we can capture signs of asymmetry. If for example, $-1 < \gamma_1 < \gamma_2 < 0$, then the negative phase of the $\hat{\varepsilon}_t$ series will tend to be more persistent than the positive phase. Enders and Siklos (2001) suggest a further alternative such that the threshold depends on the previous periods change in $\hat{\varepsilon}_t$ instead on the level of $\hat{\varepsilon}_t$. The Heaviside Indicator in this case can be set to the Momentum-Heaviside Indicator as follows:

$$I_t = \begin{cases} 1 & \text{if } \Delta\hat{\varepsilon}_{\text{t-1}} \geq \tau \\ 0 & \text{if } \Delta\hat{\varepsilon}_{\text{t-1}} < \tau \end{cases} \tag{5}$$

In this case the series $\hat{\varepsilon}_t$ exhibits more momentum in one direction than the other. The model given by (3) along with equation (5) depicts the momentum threshold autoregression (M-TAR) model. The M-TAR model can be used to capture a different type of asymmetry. If for example, $|\gamma_1| < |\gamma_2|$, the M-TAR model exhibits little adjustment for positive $\Delta\hat{\varepsilon}_{t-1}$ but substantial decay for negative $\Delta\hat{\varepsilon}_{t-1}$. Alternatively, increases tend to persist, but decreases tend to revert quickly back to the attractor irrespective of where disequilibrium is relative to the attractor. The threshold is estimated using Chan's methodology as before.[4]

To implement in this test the case of the TAR or M-TAR adjustment the Heaviside Indicator function is set according to equation (4) or equation (5) respectively and estimate equation (3) accordingly. The Φ-statistic for the null hypothesis of stationarity of $\hat{\varepsilon}_t$, i.e. $H_0:(\gamma_1=\gamma_2=0)$ is recorded. The value of the Φ-statistic is compared to the critical values computed by Enders and Siklos (2001). If we can reject the null hypothesis, it is possible to test for asymmetric adjustment since γ_1 and γ_2 converge to multivariate normal distributions (Tong 1990). The F statistic is used to test for the null hypothesis of symmetric adjustment, that is, $H_0:(\gamma_1=\gamma_2)$.

3. Empirical Analysis

The data used for this analysis are monthly average export price quotations (FOB) from July 1996 to March 2003.[5] The prices used in this study include the Ukrainian feed wheat exports and the U.S. No. 3 (Yellow) Maize. The data was obtained from the *World Grain Statistics* published by the International Grains Council. All prices are quoted in U.S. dollars. The subsequent analysis of the data is carried out on the logarithm of prices. Figure 1 illustrates the export prices of the U.S. and Ukraine.

The prices were initially tested for their order of integration using the traditional ADF test, the more powerful Elliot, Rothenberg and Stock test and the KPSS test as a confirmatory test. Table 1 below presents the results of the unit root tests for each of the price series. The unit root test results for the variables conclude that the prices are non-stationary I(1).

The results of the Engle Granger test are shown in second column of Table 2. The key point to note is that the ADF t-statistic is -2.87 indicating that the null of no cointegration cannot be rejected, which implies that both the prices are not cointegrated.

Given that the Engle Granger test has a low power to reject the null when the underlying process of adjustment is asymmetric (Enders and Granger 1998), the residuals of (1) are then estimated in the form of the TAR and M-TAR models. The results of the TAR model are shown in the third column of Table 2. The point estimates are calculated to be $\gamma_1=-0.077$ and $\gamma_2=-0.122$ which have the correct signs for convergence. The statistic $\Phi=4.30$ is lower than the 5% critical value implying that the null hypothesis of no cointegration cannot be rejected. The last column of Table 2 reports results using a consistent M-TAR model. The point estimates are found to be $\gamma_1=-0.001$ and -0.147 suggesting convergence. The statistic $\Phi=6.12$ allows us to reject the null hypothesis of no cointegration at the 10% significance level. The null hypothesis of symmetric adjustment provides a sample value of $F=3.69$ with a p-value of 0.05 implying that we can reject the null of symmetric adjustment. Given that we find asymmetric adjustment, the power of the Φ statistic in this case exceeds that of the Dickey Fuller test (Enders 2001). Finally, the Ljung Box Q statistic shows that none of the models suffer from problems of serial correlation.

The estimates of γ_1 and γ_2 are expected to be negative, suggesting convergence for the M-TAR model. Since $|\gamma_1| < |\gamma_2|$, the M-TAR model exhibits little adjustment for positive $\Delta\varepsilon_{t-1}$ but substantial decay for negative $\Delta\varepsilon_{t-1}$. In other words, increases tend to persist, but decreases tend to revert quickly back to the attractor.

4. Conclusion

The results of the Engle Granger procedure indicate that no long run relationship exists between the U.S. and Ukrainian prices. One may argue that the lack of integration of the Ukrainian wheat prices with the U.S. maize prices may be due to the high transactions costs which separates smaller markets from the world market. However, as argued above, the finding of no cointegration between the two prices may be due to the underlying process of price adjustment being asymmetric. However, where evidence of asymmetry is detected, the consistent M-TAR model best fits the data. From the M-TAR model, we find slow amounts of adjustment towards equilibrium for a positive change in the equilibrium relationship, but a substantial amount of adjustment for a negative change in the equilibrium relationship. In other words, positive changes tend to persist, while negative changes tend to revert quickly back to the attractor. Given that Ukrainian and U.S. prices co-move, when the Ukrainian price is increasing, the U.S price increases at a faster rate and when the Ukrainian price is decreasing, the U.S. price decreases at a relatively slower rate. Given that Ukraine is a minor exporter in relation to the U.S., Ukrainian prices would have to decrease at a faster rate in order to retain their market share. The fact that there is some evidence of asymmetric price adjustment is not surprising given the policy intervention that exists in the Ukrainian wheat market. The nature of price dynamics sheds light on how Ukraine has been strategically responding to U.S. maize exports since the mid 1990s. Further research is required to study the effect of national policies on price relationships.

Endnote

1. The Enders Siklos (2001) method is preferred over the Hansen and Seo (2002) test as the former allows for asymmetry based on direction, thereby allowing us to investigate the asymmetry that may exist in terms of the *momentum* of price adjustment.

2. If ω_t is not white noise, an Augmented Dickey Fuller (ADF) test may be used where lagged values of $\Delta\hat{\varepsilon}_t$ may be added to (2).

3. Chan's methodology requires the estimated residual series to be sorted in ascending order, and then eliminate the largest and smallest 15% of the $\hat{\varepsilon}_t$ series. The remaining 70% of the values were considered as possible thresholds. For each of the possible thresholds the equation was estimated using (3) and (4). The estimated threshold yielding the lowest residual sum of squares was deemed to be the appropriate estimate of the threshold.

4. In this case the same procedure is applied as before except that the 'difference' of the estimated residual was used.

5. There are a significant number of missing observations from April 2003 onwards, which is why the sample ends at March 2003.

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Table 1: Unit Root Test Results

	ERS	KPSS	Ng-Perron
UKR	-0.84 (1)	0.53** (6)	-0.81 (1)
USM	-0.50 (0)	0.71** (6)	-0.30 (0)

**Denotes significance at the 5% level. Numbers in parentheses denote the lag length or the bandwidth.

Table 2: Cointegration Results

Coefficients/Hypothesis	Engle Granger	TAR	M-TAR
γ_1	-0.10 (-2.87)	-0.077 (1.40)	-0.001 (0.025)
γ_2	N/A	-0.122 (2.59)	-0.147 (3.499)
$H_0:[\text{No Cointegration}]$	-2.87	4.30	6.12*
$H_0:[\text{Symmetry}]$	N/A	0.40 [0.52]	3.69 [0.05]
Q	11.8 [0.88]	1.01 [0.90]	0.55 [0.96]

Note: the values corresponding to Φ are compared with the Φ tables computed by Enders and Siklos (2001). *Denotes significance at the 10% level. The numbers in parentheses denote t-values. For Null hypothesis of symmetry and the Q statistics in the last two rows, the numbers in the square brackets denote p-values.

Figure 1: Export Prices (US $/ton)