Title: International Wheat Price Transmission and CAP Reform

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Introduction

The linkage between domestic and international markets can be described by the elasticity of price transmisson. The price transmission elasticity is a measure of the comovement of prices and shows the extent to which changes in world prices are transmitted back to within-country prices. It is a measure of economic market integration. Perfect price transparency will be observed among markets that are fully integrated and well functioning. Government interventions that affect imports and exports, however, will exert a downward pressure on the transmission of world prices to domestic markets.[1] On the other hand, trade liberalization will contribute to greater price transmission elasticities as domestic markets become better integrated into the world economy. As a result, price instability in world markets will be lessened as world supply and demand functions become more elastic (Bale and Lutz, 1979; Johnson, 1998). Although the commodity trade literature emphasizes the importance and usefulness of these elasticity estimates, little attention has been directed toward their estimation (Bredahl et al., 1998; Zwart and Meilke, 1979; Tyers and Anderson, 1992; Mundlak and Larson, 1992). Even less attention has been given to how elasticities change as policy reforms are implemented.

The purpose of this study is to obtain reliable estimates of the international wheat price transmission elasticity for Germany. We examine a model of cointegration over a period of policy regime changes. Specifically, we apply the threshold cointegration model of Enders and Siklos (1999) which allows us to more accurately discern the existence of a long-run equilibrium relationship among the stochastic processes as well as test for price asymmetry.

We focus attention on the estimation of the transmission elasticities and how they are impacted by policy changes. We illustrate the effects of reforms in the European Union's Common Agricultural Policy (CAP) on producer wheat prices in Germany. Of particular interest is the impact of the 1992 CAP reform and Uruguay Round Agreement on Agriculture (URAA).

The Analytical Framework

According to the law of one price, at a given point in time, the domestic commodity price should equal its world price, after adjustments are made for exchange rates, policy effects, transportation and other transactions costs, and product quality differences (Mundlak and Larson, 1992). Hence, for a particular commodity the long-run equilibrium relation is postulated:

$$P^{\mathrm{d}} = P^{\mathrm{US}} \cdot U, \tag{1}$$

where P^{d} is the domestic market price, P^{US} is the world price (in \$US) and U is the currency exchange rate.[2] All prices and exchange rates are in nominal terms. To test this relationship, we define the world price in local currency as $P^{\mathrm{w}} = P^{\mathrm{US}} \cdot U$ and rewrite (1) with lower case letters as log-transformed data:

$$p_t^{d} = \beta_0 + \beta_1 p_t^{w} + u_t \tag{2}$$

In an efficient and undistorted market $\beta_0 = 0$ and $\beta_1 = 1.0$, where u_t represents deviations from the long-run equilibrium at any period. This condition, $\beta_0 = 0$ and $\beta_1 = 1.0$, can also describe

markets that are perfectly integrated (Fackler and Goodwin, 1999). The estimate of β_1 is the price transmission elasticity. It is defined as the percentage change in the domestic price in response to a one-percent change in the world price. Tyers and Anderson (1992) refer to these elasticities as "price policy parameters" which measure the degree of market insulation, or the extent to which border prices are transmitted to the domestic market. Similarly, Duttan and Grennes (1988) interpret the transmission elasticity as a summary measure of all government policies that separate foreign and domestic markets; they capture the effect of policy on the domestic response to external changes. This "policy parameter" can be used to evaluate the impact of domestic trade distorting policies on world markets. For instance, the elasticity of net import demand or import supply is adjusted by this transmission elasticity (Bredahl et al., 1979).

Empirical Illustration

The Data

Over 22 years of monthly domestic and world (border) price data (1976:6 to 1998:12) were used to estimate the price transmission elasticities. Domestic producer (selling) prices (DM) for soft wheat in Germany were obtained from the CRONOS data bank of EUROSTAT. World prices are for No. 2 dark northern spring wheat, CIF Rotterdam ($US), were obtained from various US Department of Agriculture publications.[3] Monthly DM/$US currency exchange rates from the IMF were used to place the world price on a local currency basis.[4]

The nominal domestic and world price series were converted to logarithms and are shown in Figure 1. The data patterns reflect world supply and demand conditions, transfer and

other transaction costs and, EU agricultural policy. Abstracting from the transactions cost issue, the relationship between these two curves is largely policy induced by the CAP. But the CAP has not remained constant over this period; it has changed considerably. Hence, proper account must be taken of the policy regime changes to obtain reliable estimates of the relationship between the two stochastic processes. However, before we discuss our estimation procedures and results, we review the evolution of CAP policy and identify policy regime subperiods. The empirical analysis that follows will take explicit account of the impact of these regime periods on the international transmission of price to German wheat producers.

Figure 1 here

EU Policy Environment

For our sample period, three policy regime periods are identified. Subperiod I covers the sample from 1976:6 to 1988:6, Subperiod II from 1988:7 to 1993:6 and Subperiod III from 1993:7 to 1998:12.

Subperiod I. The CAP policy regime during this period can be characterized as the "old CAP"; a highly managed market system of administratively determined prices and protectionist policies. Because during the sample period the EU was surplus in wheat, intervention buying was the key mechanism to support internal prices above world levels. Specifically, in order to keep internal market prices from falling below the administratively set intervention price (set well above world price levels), intervention agencies would buy surpluses at the intervention price, store it and then sell on the world market at a loss or, more commonly, provide private exporters a subsidy (restitution) equal to the difference between the intervention price and the world price. The objective was to ensure that EU prices

remained in excess of, and more stable than world prices. Over time surpluses grew as farmers responded to the high internal prices, and surplus disposal costs escalated.

Subperiod II. Although public debate to reduce intervention prices intensified toward the end of 1984 and thereafter, the first major policy reform for cereals did not occur until 1988. The 1988 CAP reforms became effective July 1988. At this time, co-responsibility levies (deductions from farmers to pay for the cost of surplus production), stabilizers (increase in co-responsibility and reduction in the intervention price the following year, if production exceeded a maximum guaranteed quantity) and, voluntary set asides (with direct payments) applied to cereals. The adoption of this "stabilizer package" was touted as the start of a new era of CAP reform. Indeed, it was an effort to link price levels to output. However, these efforts were not designed to open domestic markets to the world, but rather to reduce budgetary costs via lower support prices and other cost reducing measures. While real support prices fell, the fundamental nature of the CAP remained.

Subperiod III. The CAP reform of 1992 (MacSharry) has been called the first major structural adjustment in European agricultural policy. The changes were considered so significant to warrant the name the "new CAP" (Swinbank, 1997). Although truly significant changes occurred, they were implemented within the existing CAP structure of variable levies, export restitutions and the like. This structure continued to isolate European agriculture from the world economy. Implemented in July 1993, the MacSharry reforms called for compensatory payments to farmers and a continuing lowering of prices supports to levels closer to expected world prices. The three major components of this reform were: (1) a substantial cut in intervention prices (30%), phased in over a three year period, (2) compensation to farmers for the price cuts through subsidies per hectare (area premiums), and (3) land "set aside" requirements with financial compensation for compliance; preference was

given to small farmers who could still receive payments without the set aside requirement. Even though the compensatory payments were not really decoupled form cropped area, this was a major step toward a market oriented grain economy. It was a regime change financially as well; a move from largely consumer financed (through higher prices) to where taxpayers pay a larger share (compensatory transfers). Notwithstanding the significance of these changes, the old variable levy and export subsidy structure continued to insulate the EU from world markets.

Subperiod III also includes the Uruguay Round Agreement on Agriculture (URAA). The old system of threshold prices and variable levies was abolished under the process of tariffication; these and other non-tariff barriers were converted to conventional tariffs and reduced over time. The first of the tariff cuts took place in July 1995 and the new arrangement limited the import tax so that the landed price could not exceed 155% of the intervention price. For wheat, the tariff equivalent was to be reduced 36% over a six year period. Constraints on the total level of support provided by the CAP were also imposed. It is most important to note that, unlike variable levies, with fixed import levies the landed price will rise and fall reflecting movements in the world price. This is indeed structural reform. However, with the "intervention price plus 55%" rule, a *de facto* variable levy type system remains operable at "high" reference (world price at Rotterdam) prices. Only at "low" reference prices do the fixed tariff equivalent rates apply and yield a landed price that varies directly with the world price. Continued lowering of the intervention price will result in a broader range over which domestic prices will reflect world market conditions. Until that time, however, the degree of international price transmission will remain much lower than desired i.e., the price transmission elasticity will be considerably less than 1.0. The change to tariff equivalents and limits on the volume of subsidized exports and the total level of expenditure on export subsidies became operational on July 1, 1995.

Measures of Price Variability

Measures of the variability of German and world prices were computed for each period and shown in Table 1. Over the total sample period, the variability of world prices was considerably greater than the variability of German domestic prices. This relationship is also true for each subperiod, where world price variability in subperiods I and II are some three-fold the domestic price. Even with the CAP reforms of 1992 and the subsequent URAA, the variability of domestic price in subperiod III remains less than half of that for the world. As noted above, except at very low world prices, a variable levy system remains under URAA. It is interesting to note that world prices have become progressively more stable over the three subperiods. However, the EU reforms of 1992 resulted in a noticeable increase in domestic price volatility while at the same time world price volatility lessened. For sure, a contributing factor to this "inverse relationship" in price volatility has been the worldwide trade liberalizing efforts over the past decade. For the post-URAA period by itself, the German price variability measure fell slightly (over subperiod III) while the variability of world price fell dramatically to just 1.86 percent. Indeed, monthly world wheat prices have become more stable.

Table 1 here

Integration, Cointegration and Asymmetry Tests

As a first step of the empirical investigation we implement the well-known augmented Dickey-Fuller tests (Dickey and Fuller, 1979, 1981) which consists of estimating the following regression:

$$\Delta Y_t = \alpha_0 + \alpha_1 t + \alpha_3 Y_{t-1} + \sum_{i=1}^{n} \gamma_i \Delta Y_{t-i} + \varepsilon_t \qquad (3)$$

with and without the linear time trend t. Δ denotes the difference operator $\Delta Y_t = (1 - B)Y_t = Y_t - Y_{t-1}$ and Y_t the time series under investigation. The t-statistic $\hat{t}_\mu$ tests the null hypothesis of a unit root with only a constant term in the alternative hypothesis and the t-statistic $\hat{t}_\tau$ refers to the test with a constant term and a linear time trend in the alternative hypothesis. The choice of the lag length n in equation (3) consists of estimating the number of autoregressive terms according to the largest lag with a statistically significant coefficient at the 5% level (Hall, 1994). The general-to-specific model selection procedure starts with a lag length of 16 months.

The results of the unit root tests are outlined in Table 2. According to the t-statistics $\hat{t}_\mu$ and $\hat{t}_\tau$ we cannot reject the null hypothesis of a unit root in both time series p_t^d and p_t^w. This finding is insensitive to the specification of the deterministic component and holds for the whole sample period 76:6 – 98:12 as well as for subsample periods 76:6 – 88:6, 88:7 – 93:6 and 93:7 – 98:12.[5]

Table 2 here

To estimate and test for cointegration we use the technique suggested by Enders and Siklos (1999) which is an extension of the Engle and Granger (1987) procedure. The first step of the Engle-Granger procedure consists of estimating the long-run relationship between the individual I(1) time series:

$$p_t^d = \beta_0 + \beta_1 p_t^w + u_t . \tag{4}$$

β_1 is the long-run parameter and u_t the disturbance term which may be serially correlated. The second step focuses on the estimate of ρ in the regression:

$$\Delta u_t = \rho u_{t-1} + \sum_{i=1}^{P} \gamma_i \Delta u_{t-i} + \varepsilon_t , \tag{5}$$

where the residuals from equation (4) are used. The rejection of the null hypothesis of no cointegration implies that the residuals u_t are stationary. Implicit in the Engle-Granger cointegration approach is the assumption that a tendency to move toward the long-run equilibrium is present *every time period*. Yet it is possible that movements toward the long-run equilibrium relationship need not occur in every period. According to the threshold cointegration technique proposed by Enders and Siklos the cointegrating relationship can be locally inactive inside a given range and then become active once the system gets too far from the equilibrium relationship. While the deviation from the equilibrium relationship u_t may locally have a unit root, globally the time series u_t is stationary (Balke and Fomby, 1997).

A formal way to quantify the adjustment process is:

$$\Delta u_t = I_t \rho_1 u_{t-1} + (1 - I_t) \rho_2 u_{t-1} + \sum_{i=1}^{P} \gamma_i \Delta u_{t-i} + \varepsilon_t , \tag{6}$$

where the I_t denotes the indicator variable:

$$I_t = \begin{cases} 1 \text{ if } u_{t-1} \geq 0 \\ 0 \text{ if } u_{t-1} < 0 \end{cases}. \tag{7}$$

The Enders-Siklos test approach consists in obtaining a F-statistic for the null hypothesis H_0: $\rho_1 = \rho_2 = 0$ and the comparison of this sample statistic with the critical value in Enders and Siklos (1999). With respect to the selection of the lag length n in equation (6) we again apply Hall's procedure. The Enders-Siklos procedure allows the modelling of short–run persistent deviations from the long-run equilibrium and an analysis of asymmetric adjustment. If we can reject the null hypothesis H_0: $\rho_1 = \rho_2 = 0$, it is possible to test for symmetric versus asymmetric adjustment to the long-run equilibrium using the usual F-statistic with the conventional F-distribution. The F-statistic tests the null hypothesis of symmetric adjustment H_0: $\rho_1 = \rho_2$ versus the alternative hypothesis of asymmetric adjustment.

Table 3 contains the results of the cointegration analysis. We wished to test for difference between the MacShary and URAA regimes but we only had 24 observations in the former period; not sufficient to warrant a separate regression. Thus, the regression for the subsample period 93:7 – 98:12 contains a dummy variable for the 93:7 – 95:6 period. The estimated parameter was positive. The second column of Table 3 reports the estimated long-run price transmission elasticity $\hat{\beta}_1$. As shown in Stock (1987), these estimates are superconsistent in case of cointegration. Unfortunately, we cannot present t-statistics because the corresponding standard errors are biased. As can be seen the estimated price transmission elasticity differs remarkably depending on the length of the sample period. Our results show significant increases in the transmission elasticities as increasingly liberalized policy reforms have been implemented. With respect to the results of the cointegration analysis the adjustment coefficients $\hat{\rho}_1$ and $\hat{\rho}_2$ have the correct negative sign suggesting convergence.

While we cannot reject the null hypothesis of no cointegration for the whole sample period, the F-statistics $\hat{F}_{COIN}$ for the subsample periods reject the null hypothesis indicating the existence of stable long-run relationship between both time series. Furthermore, we cannot reject the null hypothesis of symmetric adjustment relying on the statistics $\hat{F}_{ASY}$.[6]

Table 3 here

Conclusions

Reliable estimates of the international price transmission elasticities are important in order to accurately assess the impact of a policy regime change. In this paper, monthly data from 1976 to 1998 were used to examine the relationship between world and domestic wheat prices in the European Union. We identify key points of CAP reform and use this knowledge of structural change in our empirical estimation. Three policy regime periods are identified. Cointegration analysis is used to demonstrate the policy reform on the transmission of world market signals to German wheat producers. Further, computed measures of price variability show increased domestic price volatility and decreased world volatility in world prices as reforms are made to the CAP.

The long-run equilibrium transmission elasticities ranged from 0.18 during the highly administered period of the 70s and 80s to 0.30 during the post-URAA period. Our estimates are substantially lower than other estimates for the EU. For example, the elasticity estimate of Mundlak and Larson (1992) was 0.72 during the highly administered period of 1960–85.

While the EU is a long way from maintaining open agricultural markets, our results support the hypothesis that market liberalizing policies indeed increase price transparency.

[1] However, imposition of an ad valorem tariff can increase the price instability of the importing country which imposes the tariff (Bale and Lutz, 1979).

[2] In equation (1) we have adjusted only for exchange rates, not transportation or other transactions costs. Because of the close geographical proximity of the two price series, the omission of these costs will likely have little impact on the empirical results.

[3] The CIF Rotterdam prices do not include variable import levies of tariffs.

[4] The nominal price series were deflated by the monthly consumer prices indices for Germany and the US. Statistical tests showed no discernible difference between the deflated and undeflated series; hence, nominal prices were selected for analysis.

[5] In addition to the results contained in Table 2 we have performed augmented Dickey-Fuller tests including seasonal dummies in the regression equation (3) to take account of deterministic seasonality in both time series. As a result the findings are qualitatively the same.

[6] The Enders-Siklos procedure does not allow the investigation of structural breaks using, for example, Chow tests because the estimated standard errors are unreliable.

Table 1: Variability of German and World Monthly Wheat Prices (%)

Sample Period		German Prices	World Prices
Total:	76:6 – 98:12	1.76	3.90
Subperiod I:	76:6 – 88:6	1.20	4.30
Subperiod II:	88:7 – 93:6	0.92	2.62
Subperiod III:	93:7 – 98:12	1.16	2.49
URAA:	95:7 – 98:12	1.11	1.86

Note: The nominal monthly prices are in logarithms and the CIF Rotterdam prices are converted to German Marks (DM). The price variability is measured as trend-corrected coefficient of variation following the approach of Cuddy and Della Valle (1978).

Table 2: Results of the Unit Root Tests

Time Series	Sample Period	$\hat{t}_\mu$	n	$\hat{t}_\tau$	n
P^d	76:6 – 98:12	1.21	15	- 2.08	15
	76:6 – 88:6	0.32	12	- 0.34	12
	88:7 – 93:6	- 2.78	12	- 0.84	12
	93:7 – 98:12	- 0.05	13	- 2.84	12
P^w	76:6 – 98:12	- 1.91	7	- 2.46	7
	76:6 – 88:6	- 1.49	0	- 1.49	0
	88:7 – 93:6	- 2.49	6	- 0.96	14
	93:7 – 98:12	- 1.38	1	- 2.56	1

Note: $\hat{t}_\mu$ and $\hat{t}_\tau$ are *t*-statistics of the Dickey-Fuller tests corresponding to a regression with a constant term and, a constant term as well as a linear time trend, respectively. *n* denotes the lag length which is chosen according to the largest lag with a statistically significant coefficient at the 5% level (Hall, 1994). Critical values can be found in MacKinnon (1991).

Table 3: Results of the Cointegration Analysis

Sample Period	$\hat{\beta}_1$	$\hat{\rho}_1$	$\hat{\rho}_2$	$\hat{F}_{COIN}$	$\hat{F}_{ASY}$	n
76:6 – 98:12	0.54	– 0.03	– 0.01	1.61	0.26	5
76:6 – 88:6	0.18	– 0.16	– 0.08	5.44*	0.64	1
88:7 – 93:6	0.18	– 0.37	– 0.26	6.94*	0.19	1
93:7 – 98:12	0.25	– 0.48	– 0.26	10.14*	0.93	1
95:7 – 98:12	0.30	– 0.39	– 0.36	7.03*	0.02	1

Note: $\hat{\beta}_1$ is the estimated long-run elasticity from the regression $P^d = \beta_0 + \beta_1 Pw + u_t$, $\hat{\rho}_1$ and $\hat{\rho}_2$ the estimated adjustment coefficients, $\hat{F}_{COIN}$ the *F*-statistic to test the null hypothesis of no cointegration and $\hat{F}_{ASY}$ the *F*-statistic to test symmetric adjustment. * denotes significant test statistic at a 5% level. Critical values can be found in Enders and Siklos (1999).

References

Bale, M.D. and Lutz, E. "The Effects of Trade Intervention on International Price Instability".
Amer. J. Agr. Econ. 61 (August, 1979): 512 – 16.

Balke, N.S. and Fomby, T. "Threshold Cointegration". *Intern. Econ. Rev.* 38, (August, 1997):
627 – 45.

Bredahl, M.E., Meyers, W.H and Collins, K.J. "The Elasticity of Foreign Demand for U.S.
Agricultural Products: The Importance of the Price Transmission Elasticity". *Amer. J.
Agr. Econ.* 61 (February, 1979): 58 – 63.

Dickey, D.A. and Fuller, W.A. "Distribution of the Estimators for Autoregressive Time Series
with a Unit Root". *J. Amer. Statist. Assoc.* 74 (June, 1979): 427 – 31.

Dickey, D.A. and Fuller, W.A. "The Likelihood Ratio Statistics for Autoregressive Time
Series with Unit Roots". *Econometrica.* 49 (July, 1981): 1057 – 72.

Dutton, J and Grennes, T. "The Role of Exchange Rates in Trade Models". Chapter 4 in *Elasticities in International Agricultural Trade*, eds., Carter, C.A. and Gardiner, W.H., Westview Press 1988.

Enders, W. and Siklos, P.L. "Cointegration and Threshold Adjustment". *Working Paper* Wilfrid Laurier University Waterloo, Ontario, Canada (1999).

Engle, R.F. and Granger, C.W.J. "Co-Integration and Error Correction Representation, Estimation and Testing." *Econometrica*. 55 (March, 1987): 251 – 76.

Fackler, P.L and Goodwin, B.K. "Spatial Price Analysis". Handbook of Agricultural Economics, *Amer. Agr. Econ. Assoc.*, forthcoming, 1999.

Hall, A. "Testing for a Unit Root in Time Series with Pretest Data-Based Model Selection". *J. Bus. Econ. Stat*. 12 (October, 1994): 461 – 70.

Johnson, D.G. "Food Security and World Trade Prospects". *Amer. J. Agr. Econ*. 80 (December, 1998): 941 – 47.

MacKinnon, J.G. "Critical Values for Cointegration Tests, in Long-Run Economic Relationships". *Readings in Cointegration*, (Eds) R.F. Engle, and C.W.J. Granger, Oxford University Press, Oxford (1991): 267 – 76.

Mundlak, Y. and Larson D.F. "On the Transmission of World Agricultural Prices". *World Bank Econ. Rev*. 6 (September, 1992): 399 – 422.

Stock, J. H. "Asymptotic Propertries of Least Squares Estimators of Co-Integrating Vectors". *Econometrica* 55 (September, 1987): 1035 – 1056.

Swinbank, A. "The New CAP" in *The Common Agricultural Policy*. Eds. Ritson, C. and D.R. Harvey. (London: CAB International), 1997.

Tyers, R. and Anderson, K. *Disarray in World Food Markets*: *A Quantitative Assessment*. Cambridge: Cambridge University Press, 1992.

Zwart, A.C. and Meilke, K.D. "The Influence of Domestic Pricing Policies and Buffer Stocks on Price Stability in the World Wheat Industry". *Amer. J. Agr. Econ.* 61 (August, 1979): 434 – 44.