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REGIONAL AND INTERNATIONAL MARKET INTEGRATION OF A SMALL OPEN ECONOMY

SEBASTIAN FOSSATI
University of Washington

FERNANDO LORENZO
Centro de Investigaciones Económicas, Uruguay

CESAR M. RODRIGUEZ*
University of Pittsburgh

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This paper studies the relationship between a set of commodity prices in a small open economy like Uruguay and the corresponding international and regional prices. The empirical methodology used is the multivariate cointegration procedure based on maximum likelihood methods introduced by Johansen (1988) as well as estimations of half-life persistence indicators. In the case of cereals, the evidence suggests strong market integration between domestic and regional markets and, to some extent, also to international markets. Therefore, directly or indirectly, domestic prices are connected with the efficient price signal. Results for beef indicate strong market integration between the domestic market and the regional market, which is not so well connected with international markets. Thus, domestic price appears to be linked to a regional price that is not linked to the efficient price signal.

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I. Introduction

The relationship between prices has a long history in economics and has been

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* César M. Rodríguez (corresponding author): cmr37@pitt.edu. The authors are grateful to three anonymous referees and the co-editor for detailed comments on an earlier draft. They also thank participants at the XXI Latin American Meeting of the Econometric Society at Panama City, Panama, the Middle East Technical University VII International Conference in Economics at Ankara, Turkey, and the VIII Annual Meeting of the Latin American and Caribbean Economic Association (LACEA) at Puebla, Mexico, for comments.

used to define a market as early as the 19^{th} century. Cassel (1918) seems to be the earliest reference in relation to international trade introducing the notion of purchasing power parity and the law of one price (LOOP). More recently, Stigler (1969) defines a market as “the area within which the price of a commodity tends to uniformity, allowance being made for transportation costs”. Based on this definition, there exists a large empirical literature investigating market integration by analyzing price relationships.[1]

The LOOP states that for a given commodity, a representative price, adjusted by exchange rates and allowance for transportation costs, will prevail across all countries. Therefore, the LOOP suggests that similar commodity markets across all countries should be integrated as a single market, which is warranted by efficient international commodity arbitrage.

Geographically separated markets are spatially integrated if goods and information flow freely among them and, as a result, the effects of price changes in one market are transmitted to another market’s price. Theoretically, under the assumption of perfect competition, when two regions trade, the product price in the import region equals the price in the export region plus transportation cost. Therefore, the price change in the export region induces a price change in the import region in the same direction and of the same degree. If this is the case, the two markets are completely integrated as a single market. The extent and the speed to which shocks are passed through, and the strength of the interdependence among prices are indicators of the degree of integration and global efficiency of markets’ performance. As pointed out by Ravallion (1986), measurement of market integration can be viewed as basic data for understanding how specific markets work. The extent to which commodity markets are integrated also has important implications for governments’ regulation and general economic policy.[2]

The issue of price convergence in commodity markets both at national and international level has been studied in the literature rather extensively either under

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[1] These approaches have their deficiencies, and certainly provide less information than partial equilibrium models of markets, where demand and supply equations are specified. However, since price data is available to a much larger extent than quantity data, price analysis will be possible in many cases where analysis using other approaches is not.

[2] If the market is globally integrated, government intervention within one nation may be ineffective or very costly.

the notion of the LOOP (Ardeni 1989, Baffes 1991) or under the notion of market integration (Ravallion 1986). Recognizing the nonstationarity property of commodity prices, researchers have extensively employed cointegration and error correction models (ECM) (Engle et al. 1987) to test the LOOP and market integration on international commodity markets. This is particularly useful because the LOOP and market integration are tested as a long-run relationship that is not affected by short-run deviations. Earlier studies already found that the LOOP almost never holds in the short-run. These include Ardeni (1989), Hazel et al. (1990), Mundlak et al. (1992), Baffes (1991), Goodwin (1992), Zanias (1993), Barrett (1996), Fackler (1996), Yang et al. (2000a, 2000b), and Bukenya et al. (2002). Most of these authors found some evidence for the validity of the LOOP and international market integration.

This paper explores the relationship between a set of commodity prices (maize, wheat, sorghum and beef) in a small open economy like Uruguay and the corresponding international and regional prices. The relevant regional commodity prices for Uruguay come from Argentina, while the relevant international prices are from the United States (for maize, wheat and sorghum) and Australia-New Zealand (for beef). As Argentina is a major exporter of maize, wheat and to some extent sorghum, just like the United States, one thing worth mentioning is that the regional relevant price for Uruguay, with the exception of beef prices, is not a "small player" price. Therefore, this paper addresses not only the price relationship between a small open economy (Uruguay) and the "big players", but also between large developed countries (United States and Australia) and a large developing one (Argentina).

Two competing hypotheses are analyzed in this paper. The first one refers to the extent and the pattern of market integration among these countries and therefore is a natural outcome of the LOOP hypothesis. In most studies using cointegration tests to investigate market integration, Engle and Granger's test has been the preferred tool, although some recent exceptions exist. This test has several notable weaknesses. The most important in a market delineation context are that hypothesis on the estimated parameters cannot be tested and that the estimates of the cointegration vectors depend on the choice of the dependent variable. It is well known that these problems can be avoided by using Johansen's (1988, 1995) multivariate cointegration procedure, which in this paper is applied in the context of international market integration. The other issue tackled is the degree of integration among markets calculated with half-life persistence indicators following Taylor (2001). Thus, the degree of integration between countries that belong to the same market is defined as the reaction time to remove disequilibria after a shock.

This study contributes to the literature in the following two ways. First, it extends the work done by Yang et al. (2000a, 2000b) by examining several commodities providing a more comprehensive perspective on market integration. Second, this study also addresses both at regional and international levels the market integration of a small open economy.

The paper is organized as follows. The next section discusses the empirical methodology applied and the hypothesis testing based on the Johansen (1988) cointegration procedure. Section III describes the data. Section IV discusses the empirical results. Finally, some concluding remarks are drawn in Section V.

II. Empirical methodology and hypothesis testing

A. Cointegration analysis

Multiple techniques have been used to analyze product market integration, most of them looking at the relationship between prices of different markets. If two markets are spatially integrated then price signals in one market should reflect in the other one; so the simplest way to test spatial market integration is calculating correlation coefficients. A natural extension is regression analysis; however, since correlation analysis is static rather than dynamic, it is also important to examine cross-correlations with a lag structure between the variables of interest.

The dynamic dimension of market integration was introduced in Ravallion's (1986) seminal work providing the definition of short and long-run market integration. Nevertheless, important shortcomings from previous approaches, derived from the univariate properties of price series, have driven market integration researchers towards the so-called cointegration approach.[3] If two price series are cointegrated, they tend to move towards an equilibrium relationship in the long-run and, therefore, their respective markets are integrated. Engle and Granger's

[3] Price series used for these analyses are usually nonstationary.

(1987) bivariate procedure has been widely used for testing cointegration among agricultural price series. However, when using Engle and Granger's test, it is not possible to analyze hypotheses on the parameters (the cointegration vector). Thus, it is not possible to test the hypothesis of market integration provided by the specification given by the LOOP.

Johansen's (1988) multivariate cointegration procedure is more adequate as multilateral trade is likely to induce a simultaneous determination of market prices (i.e., exogeneity must be tested rather than assumed). Moreover, the need to account for the dynamic structure jointly with the long-term structure, the possibility of dealing simultaneously with multiple cointegration relationships and testing restrictions on the parameters in order to test the LOOP makes Johansen's procedure preferable. Hence, the procedure used in this paper is the multivariate procedure based on maximum likelihood methods introduced by Johansen (1988, 1991) and expanded upon by Johansen et al. (1990, 1994).

Johansen's (1988) procedure is based on a vector autoregressive model of X_t, a ($nx1$) vector of $I(1)$ time series.[4] The error-correction form is written in first differences as:

$$\Delta X_t = A_1 \Delta X_{t-1} + \ldots + A_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \varepsilon_t , \qquad (1)$$

$$\varepsilon_t \sim N(0, \Lambda), \qquad t = 1, \ldots, T ,$$

where A_i for all i (i=1...k-1) and Π are (nxn) matrices, μ is a ($nx1$) vector of constants, ε_t is a ($nx1$) error vector and Λ is its (nxn) covariance matrix. Since ΔX_t is an $I(0)$ process, the stationarity of the right side of the equation is achieved only if ΠX_{t-k} is stationary.

Johansen's procedure examines the rank of Π, which determines the number of cointegrating vectors present in the system. If $rank(\Pi) = r < n$, then $\Pi = \alpha\beta'$, where both α and β are (nxr) matrices. β is the matrix of cointegrating vectors, and the number of such vectors is r. Since the cointegrating vectors have the property that $\beta_j' X_t$, for all j (j=1...r) is stationary, then the system is stationary. The cointegrating vectors are said to represent the long-term relationships present in the system.

[4] In this case n = 3 [Domestic Price (*Pd*), Regional Price (*Pr*) and International Price (*Pi*)]. In addition, it should be noted that equation (1) is intended to be used for individual commodities.

The first hypothesis concerns whether there is some degree of market integration among some or all of these markets (domestic, regional, international) for each product. Long-run market integration would mean that there is a long-term relationship between prices, i.e., that there is a cointegrating vector. This hypothesis can be tested by examining the rank of Π using the trace and max statistics defined by Johansen (1988, 1991). The trace statistic tests whether r cointegrating vectors are present in the system against the alternative hypothesis that the system is already stationary (i.e., n cointegrating vectors are present in the system). Equivalently, the max statistic tests whether the rank is r against the alternative hypothesis that the rank is $(r+1)$. The distributions of the test statistics are non-standard, and approximate asymptotic critical values were tabulated by Osterwald-Lenum (1992).

It is important to include a constant term, μ, in equation (1) when calculating the test statistics in order to take transportation costs and quality price differentials into consideration. If these two effects are relatively constant in the long-run, we may restrict the constant term to the cointegration space. Otherwise, μ may capture transportation costs and quality price differentials with a linear time trend.

The second hypothesis concerns whether some of these markets are not constrained to the cointegrating relationship. The hypothesis can be tested by examining whether $\beta_{ij}=0$ for all j (j=1...r) cointegration vectors for the i^{th} market (i=1...n) using the appropriate likelihood ratio test statistics.

Among those markets that are confirmed to be integrated in the second hypothesis, the third hypothesis tests whether their prices tend to be equal, allowing for price differences due to transportation costs and quality price differentials. The hypothesis can be tested by examining whether $\beta_{ij}-\beta_{kj}=0$ for all j (j=1...r) cointegration vectors for the i^{th} and k^{th} markets (i,k=1...n) using the likelihood ratio test statistics.

Finally, weak exogeneity is tested among those markets that are confirmed to be integrated in a single market in the second and third hypotheses. If the i^{th} market price X_i is weakly exogenous, it does not respond to the deviations from the relevant long-run relationship and can be considered as one of the forces that "guides" the system. The hypothesis can be tested by examining whether $\alpha_{ij}=0$ for all j ($j=1...r$) cointegration vectors for the i^{th} market ($i=1...n$) using the appropriate likelihood ratio test statistics. The size of the adjustment coefficients (α) also provides information about the degree of market integration, however, this paper also pretends to estimate the convergence speed of prices, i.e., how persistent are price gaps, calculating half-life indicators.

When analyzing how many long-run relationships appear among the price series considered in this paper, it is possible to identify two cases: the case of two cointegrating vectors and the case of only one cointegrating vector. When two cointegrating vectors are found, the long-run relationships can be written as:

$$\alpha\beta' P_{t-1} = \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \tag{2}$$

The perfect market integration hypothesis can be tested imposing a set of restrictions on β which yields the following β^* matrix:

$$\beta^* = \begin{bmatrix} 1 & -1 & 0 & * \\ 0 & 1 & -1 & * \end{bmatrix} \tag{3}$$

where * denotes unrestricted constants in the cointegration space allowing for transport costs and quality price differentials. The second row implies that the regional price is equal to the international price, and the first row implies that the domestic price is equal to the regional price.

When there is only one cointegrating vector among the three variables, the analysis must take into consideration the possibility that one of the variables does not appear in the long-run relationship. In the case of one cointegrating vector, the long-run relationship can be written as:

$$\alpha\beta' P_{t-1} = \begin{bmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \tag{4}$$

B. Half-life indicators

In order to assess the degree of integration among markets, half-life persistence indicators are calculated. The half-life indicators are calculated in pairs of regions for each product. Compared to the speed of adjustment coefficients (α), this method has the advantage that half-life indicators can be calculated independently of the existence of a long-run relationship between the prices. However, if prices are cointegrated then low half-lives are expected but if prices are not cointegrated, then half-lives would tend to infinity.

Following Taylor (2001), the basic model of price convergence considers a linear AR(1):

$$x_t = \rho \cdot x_{t-1} + \varepsilon_t \tag{5}$$

where $x_t = p_t^2 - p_t^1$ is the price gap between two markets measured in a common currency, ε_t is $N(0,\sigma^2)$, $0 < \rho < 1$ and $\lambda = \rho - 1 < 0$ is the convergence speed. The half-life of the deviations, i.e., how much time (measured in months) is needed so that 50% of the total long-run effect of the shock is transmitted, is:

$$H = \frac{ln(0.5)}{ln(\rho)} \tag{6}$$

If price gaps are stationary ($0 < \rho < 1$) then shocks will have temporary effects and so the half-life of the deviations will be small. But, if price gaps are nonstationary (they might follow a random walk, $\rho = 1$) then shocks will have permanent effects and so half-life will be infinite.

III. Data

Commodity price time series data for maize, wheat, sorghum and beef, covering two decades for almost all products, are utilized in this paper. For maize, the data covers the period from 1981.04 to 2000.12; for wheat from 1974.04 to 2000.12; for sorghum from 1981.04 to 2000.12 and finally for beef from 1981.04 to 2000.12. The representative international and regional price signal for a small open economy like Uruguay depends on the commodity under consideration.

The data used in this paper includes domestic, regional –taking Argentina as a reference– and international –United States (Gulf Ports) for maize, wheat and sorghum and Australia-New Zealand for beef– prices. Domestic maize, wheat, sorghum and beef prices were obtained from OPYPA while the relevant regional prices from Argentina were obtained from the Direccción de Mercados Agroalimentarios, SAGPyA.[5] Finally, relevant international prices were obtained from the International Monetary Fund's International Financial Statistics (version 1.1.53). Price differences due to quality differences as well as transportation costs may be captured by a properly defined constant term in the cointegration model, as explained in the previous section.

Two features of the data set are worth mentioning. First, results of this study are more likely to be free from the influence of governmental price controls. It has been argued that government intervention can fundamentally change cointegration of international commodity prices (Bessler et al. 1996, Yang et al. 1999). There are no direct government price controls that affect maize, wheat, sorghum or beef prices, thus, these prices can be significantly more market-driven than many other agricultural commodities.

Second, at least in maize, wheat and to some extent sorghum, Argentina is a major exporter of these commodities just like the United States. The latter accounts for the fact that the regional relevant price for Uruguay is not a small player price. The exception is beef where Argentina is more like a typical small economy. Theoretical models of open economies typically suggest that small open economies are much more likely to follow the prices determined by the big players (usually the large developed countries) in international commodity markets, whether they are developed or still developing. Therefore, this paper addresses not only the price relationship between a small open economy (Uruguay) and the big players, but also between large developed countries (United States and Australia) and a large developing one (Argentina).

IV. Empirical results

When investigating for market integration, the first step consists in examining each price series for evidence of nonstationarity in order to confirm that the cointegration approach is appropriate. This analysis was performed using the

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[5] OPYPA: Oficina de Programación y Política Agropecuaria, from the Ministerio de Ganadería, Agricultura y Pesca, Uruguay; SAGPyA: Secretaría de Agricultura, Ganadería, Pesca y Alimentación, Argentina.

Augmented Dickey-Fuller test which allowed us to consider the variables as integrated of order one, *I(1)*.[6] The next step would be to perform the cointegration tests.

A. Sorghum

The analysis starts considering the case of sorghum, where the first step is to determine the lag length for the vector autoregression.[7] Simplification tests of the initial system lag length (fifteen) suggested that eight lags sufficed, so this reduction was implemented. Table 1 reports the results of the rank tests, which suggest that there are two cointegrating vectors among the three markets considered. According to Stock et al. (1988) these three variables have a common trend, thus, there is some kind of market integration among them.

Of the two long-run relationships identified, one cointegrating vector may represent the existence of a long-run relationship between the two major markets, Argentina and US Gulf Ports, in the international market. The other cointegrating vector may represent a long-run relationship between the domestic market and the international markets. The next hypothesis is whether market integration found among these markets is "perfect" or not, which implies that prices should be equal allowing for differences due to transportation costs and quality price differentials.

The likelihood ratio test results are summarized in panel C of Table 1. The χ^2 test statistics suggest no rejection of the projected restrictions at 0.05 level (p-value 0.126). Several alternative restrictions in different combinations were tried, but all these identification hypotheses were rejected. Consequently, the results suggest that the LOOP structure, in its relative version, cannot be rejected as a single price holds across the three countries considered.

Finally, the dynamic adjustment to the long-run common trend must be determined by performing weak exogeneity tests, reported in panel C of Table 1. Weak exogeneity of the international price cannot be rejected and so the US Gulf Ports are the primary source of information that drives the single common trend in the long-run. Combining the identified LOOP structure in β and the α matrix led to the following specification of system (2) (p-value 0,112):

[6] Fossati and Rodríguez (2002) analyzed the univariate properties of the price series considered in this paper, performing unit root tests, seasonal unit root tests and univariate modelling, concluding that the series are integrated of order one. For reasons of space, the results are not presented, but are available upon request.

[7] The tests were performed using PcFiml 9.00.

Table 1. Sorghum (1981:04 - 2000:12)

Panel A: Rank test on Π				
H_0: rank = r	Q_{max}	95%[1]	Q_{trace}	95%[1]
r == 0	28.06**	22.0	57.68**	34.9
r <= 1	20.62**	15.7	29.63**	20.0
r <= 2	9.01	9.2	9.01	9.2

Panel B: Model evaluation				
Statistics	*Pd*	*Pr*	*Pi*	VECM
$F_{ar}(7, 205)$	0.94	1.35	1.29	-
$F_{arch}(7, 198)$	1.57	0.09	0.11	-
$F_{het}(48, 163)$	0.75	1.21	1.19	-
$\chi^2_{nor}(2)$	64.97**	252.83**	164.97**	-
$F_{ar}(63, 565)$	-	-	-	1.10
$F_{het}(288, 950)$	-	-	-	1.08
$\chi^2_{nor}(6)$	-	-	-	397.64**

Panel C: Likelihood ratio tests results		
Hypothesis	χ^2 statistics	Degrees of freedom
H_1: $\beta_{11} = \beta_{22} = 1; \beta_{21} = \beta_{32} = -1; \beta_{31} = \beta_{12} = 0;$	4.1403	2
H_2: $\alpha_{11} = \alpha_{12} = 0$	18.203 **	2
H_3: $\alpha_{21} = \alpha_{22} = 0$	8.5335 *	2
H_4: $\alpha_{31} = \alpha_{32} = 0$	0.5303	2
H_5: $\alpha_{12} = \alpha_{21} = \alpha_{31} = \alpha_{32} = 0$	0.5303	2
H_6: $H_1 + H_5$	12.771 *	6
H_7: $\beta_{11} = \beta_{22} = 1; \beta_{21} = \beta_{32} = -1; \beta_{31} = \beta_{12} = 0;$ $a_{21} = \alpha_{31} = \alpha_{32} = 0;$	8.9329	5

Panel D: Half-lives and AR(1) coefficients

	Pd – Pr	*Pd – Pi*	*Pr – Pi*
Half-life	4.77	3.99	7.36
AR(1)	0.865	0.841	0.910

[1] Critical values from Osterwald-Lenum (1992). *(**) denotes significance at the 5% (1%) level.

$$\alpha\beta' P_{t-1} = \begin{bmatrix} -0.18 & -0.25 \\ 0 & -0.35 \\ 0 & 0 \end{bmatrix} \begin{bmatrix} 1 & -1 & 0 & -0.15 \\ 0 & 1 & -1 & 0.13 \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \tag{7}$$

Results show a significant negative response of the domestic price to perturbations in all the two cointegrating vectors. When the domestic price is high relative to its long-run relationship with the regional price, the domestic price falls in the next period. In a similar way, when the regional price is high relative to its long-run equilibrium with the international price, the domestic price falls in the subsequent period. Therefore, the domestic price follows both movements in the regional and international prices in the same direction in the subsequent period. In the case of the regional price, when it is high to its relative long-run equilibrium with the international price, it falls in the subsequent period.

These results imply that, in the case of sorghum, the US Gulf Ports market has the price leadership, while the markets in Uruguay and Argentina are price followers. Results also show a larger initial response in Argentina's prices (0.35) than in Uruguay's prices (0.25) suggesting that Argentina's sorghum market is more integrated to international markets than Uruguay's market.

Panel D of Table 1 reports the estimated half-life of deviations (measured in months) for each pair of prices. Estimated half-lives, less than eight months in all cases, imply an elevated grade of market integration between these markets as expected from the cointegration analysis results. Still, the lower speed of adjustment between the regional and international market was not expected. However, considering the cointegration analysis results, it would be necessary to analyze if these differences between estimated half-life coefficients are significant.

B. Maize

The next product to consider is maize. The lag length for the vector autoregression is four. Table 2 reports the results of the rank tests, which suggest that there are two cointegrating vectors among the three markets considered (US Gulf Ports, Argentina, Uruguay) implying that there is a long-run common trend among them.

The perfect market integration hypotheses among these markets are presented in panel C of Table 2 where the likelihood ratio test results are summarized. The χ^2 test statistics for (3) rejects the projected restrictions at 0.05 level (p-value 0.008). Relaxing the restrictions in different combinations led to the identification of β matrix. The weak exogeneity tests, summarized in panel C of Table 2, suggest that weak exogeneity is rejected for all markets. Combining the identified structure on β and the α matrix led to the following specification of system (2) (p-value 0.845):

$$\alpha\beta' P_{t-1} = \begin{bmatrix} -0.22 & 0 \\ 0 & -0.32 \\ 0 & 0.17 \end{bmatrix} \begin{bmatrix} 1 & -0.41 & 0 & -5.14 \\ 0 & 1 & -1 & 0 \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \qquad (8)$$

In the case of maize, the LOOP structure in its absolute version is not rejected for the regional price and the international price. Moreover, the weak exogeneity tests suggest that neither the international price nor the regional price is the primary source of information that drives the single common trend in the long-run. These results imply that these two markets share the price leadership and both respond to perturbations to its long-run relationship. Specifically, the response of the regional price (0.32) is larger than the response of the international price (0.17), giving evidence of the relative importance (share) of Argentina in the world maize market.

In the first cointegrating vector, the hypothesis of perfect market integration is rejected. In this case, when the domestic price is high relative to its long-run relationship with international prices, there is a significant negative response of the domestic price in subsequent periods. The α matrix suggests that the market of maize in Argentina could be isolating the domestic market (Uruguay) from the international markets as there is no direct relationship between the domestic price and the international price. Moreover, price signals from international markets seem to be transmitted to the domestic market through Argentina.

When considering the results from the half-life indicators in panel D of Table 2, further evidence on the pattern of market integration was found. Specifically, price signals from international markets are transmitted at a higher speed to the regional market than to the domestic market, suggesting a larger degree of market integration of the regional market. In the case of the domestic market, results suggest a stronger

Table 2. Maize (1981:04 - 2000:12)

Panel A: Rank test on Π				
H_0: rank = r	Q_{max}	95%[1]	Q_{trace}	95%[1]
r == 0	61.19**	22.0	101.5**	34.9
r <= 1	31.14**	15.7	40.32**	20.0
r <= 2	9.18	9.2	9.18	9.2

Panel B: Model evaluation				
Statistics	*Pd*	*Pr*	*Pi*	VECM
$F_{ar}(7, 217)$	1.40	0.62	1.08	-
$F_{arch}(7, 210)$	1.63	0.29	0.28	-
$F_{het}(24, 199)$	1.08	0.62	1.30	-
$\chi^2_{nor}(2)$	92.07**	75.86**	144.88**	-
$F_{ar}(63, 600)$	-	-	-	1.23
$F_{het}(144, 1141)$	-	-	-	1.59**
$\chi^2_{nor}(6)$	-	-	-	209.55**

Panel C: Likelihood ratio tests results Hypothesis	χ^2 statistics	Degrees of freedom
H_1: $\beta_{11} = \beta_{22} = 1; \beta_{21} = \beta_{32} = -1; \beta_{31} = \beta_{12} = 0;$	9.6337**	2
H_2: $\beta_{11} = \beta_{22} = 1; \beta_{32} = -1; \beta_{31} = \beta_{12} = \beta_{42} = 0;$	1.923	2
H_3: $\alpha_{11} = \alpha_{12} = 0$	21.838**	2
H_4: $\alpha_{21} = \alpha_{22} = 0$	19.442**	2
H_5: $\alpha_{31} = \alpha_{32} = 0$	7.4769*	2
H_6: $\alpha_{12} = \alpha_{21} = \alpha_{31} = 0$	0.0001	1
H_7: $H_2 + H_6$	2.0283	5

Panel D: Half-lives and AR(1) coefficients	*Pd – Pr*	*Pd – Pi*	*Pr – Pi*
Half-life	9.38	11.06	2.70
AR(1)	0.928	0.939	0.774

[1] Critical values from Osterwald-Lenum (1992). *(**) denotes significance at the 5% (1%) level.

degree of market integration between the domestic price and the regional price than between the domestic price and the international price as price signals are transmitted at a higher speed from regional than from international markets.

C. Wheat

In the case of wheat, the lag length for the vector autoregression determined by the F-test was seven. Table 3 reports the results of the rank tests, suggesting there is one cointegrating vector among the three markets considered (US Gulf Ports, Argentina, Uruguay) which implies that there are two long-run trends among them.

The likelihood ratio test results of the exclusion, exogeneity and homogeneity tests for wheat are summarized in panel C of Table 3. Combining the restrictions on β and α led to the following specification of system (4) (p-value 0.106):

$$\alpha\beta' P_{t-1} = \begin{bmatrix} 0 \\ -0.16 \\ 0 \end{bmatrix} \begin{bmatrix} 0 & 1 & -1 & 0.08 \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \tag{9}$$

The cointegrating vector found relates the regional price to the international price in the same fashion as in the case of sorghum. This result suggests that LOOP structure, in its relative version, cannot be rejected as a single price holds across the regional and international markets. However, the speed of adjustment of the regional price to deviations from its long-run relationship with the international price (0.16) is lower than in the cases of sorghum or maize.

In this case, results suggest that the domestic wheat market is not integrated to the international markets since no cointegrating vector that relates the domestic price to the regional and/or the international prices was found. This result is quite disturbing as there are no big differences between the market structure of wheat and the other cereals (Fossati and Rodriguez 2002). However, the fact that domestic price series are very noisy could explain the lack of a stationary long-run relationship between the domestic price and both the regional and the international prices (Fossati and Rodriguez 2002).

Results from the half-life indicators in panel D of Table 3 show that the 50% of

Table 3. Wheat (1974:04 - 2000:12)

Panel A: Rank test on Π

H_0: rank = r	Q_{max}	95%[1]	Q_{trace}	95%[1]
r == 0	25.93 *	22.0	44.04 **	34.9
r <= 1	10.56	15.7	18.11	20.0
r <= 2	7.54	9.2	7.54	9.2

Panel B: Model evaluation

Statistics	*Pd*	*Pr*	*Pi*	VECM
$F_{ar}(7, 292)$	2.58 *	1.29	2.14 *	-
$F_{arch}(7, 285)$	0.08	0.75	2.60 *	-
$F_{het}(42, 256)$	1.62*	0.86	1.03	-
$\chi^2_{nor}(2)$	185.7**	171.9**	25.1**	-
$F_{ar}(63, 824)$	-	-	-	1.71 **
$F_{het}(252, 1501)$	-	-	-	1.32 **
$\chi^2_{nor}(6)$	-	-	-	403.73 **

Panel C: Likelihood ratio tests results

Hypothesis	χ^2 statistics	Degrees of freedom
H_1: $\beta_{11} = 0$	1.4256	1
H_2: $\beta_{21} = 0$	13.144 **	1
H_3: $\beta_{31} = 0$	15.258 **	1
H_4: $\beta_{41} = 0$	5.9162 *	1
H_5: $\alpha_{11} = 0$	0.005	1
H_6: $\alpha_{21} = 0$	7.6283 *	1
H_7: $\alpha_{31} = 0$	2.5331	1
H_8: $H_1 + H_5 + H_7$	3.0174	3
H_9: $\beta_{11} = 0; \beta_{21} = 1; \beta_{31} = -1; \alpha_{11} = \alpha_{31} = 0$	7.6458	4

Panel D: Half-lives and AR(1) coefficients	*Pd – Pr*	*Pd – Pi*	*Pr – Pi*
Half-life	5.25	5.85	3.75
AR(1)	0.876	0.888	0.831

[1] Critical values from Osterwald-Lenum (1992). *(**) denotes significance at the 5% (1%) level.

a price shock is transmitted in less than a year, showing the elevated grade of market integration. Moreover, the degree of market integration between the regional and international markets is considerably large. It is important to note that under this methodology there is no difference between wheat and the other cereals, as expected. Therefore, the domestic price shows a stronger degree of market integration with the regional price than with the international price as price signals are transmitted at a higher speed from regional than from international markets.

D. Beef

The lag length selected by the F-test for the vector autorregression for beef was four. Table 4 reports the results of the rank tests, which suggest that there is one cointegrating vector among the three markets considered (AU-NZ, Argentina, Uruguay) implying that there are two long-run trends among them.

In the case of beef, the products that are being studied are not entirely homogeneous, so the LOOP structure hypothesis makes no sense. The likelihood ratio test results are summarized in panel C of Table 4. Combining the restrictions on β and α led to the following specification of system (4) (p-value 0.086):

$$\alpha\beta' P_{t-1} = \begin{bmatrix} -0.14 \\ 0 \\ 0 \end{bmatrix} \begin{bmatrix} 1 & -0.42 & 0.54 & 0.67 \end{bmatrix} \begin{bmatrix} Pd_{t-1} \\ Pr_{t-1} \\ Pi_{t-1} \\ 1 \end{bmatrix}. \tag{10}$$

The international price (AU-NZ in this case) is found to be weakly exogenous to the system, showing the importance of that market in the international beef sector. Still, the regional price is weakly exogenous for this system too. This result can be explained by the fact that, in the sample period, more than 90% of the beef production in Argentina is addressed to domestic consumption.

The domestic price appears to be connected with both the regional price and the international price, however, the sign of the coefficient for the AU-NZ price is not the expected. In particular, the domestic price is the one adjusting to disequilibria in the long-run relationship. The fact that the domestic price is linked to the regional price, which is not linked to the international price, is one of the outstanding results in the analysis of the beef sector.

Table 4. Beef (1980:01 - 2000:12)

Panel A: Rank test on Π				
H_0: rank = r	Q_{max}	95%[1]	Q_{trace}	95%[1]
r == 0	33.93*	22.0	39.99*	34.9
r <= 1	4.21	15.7	6.07	20.0
r <= 2	1.85	9.2	1.85	9.2

Panel B: Model evaluation				
Statistics	*Pd*	*Pr*	*Pi*	VECM
$F_{ar}(7, 232)$	0.73	0.84	0.98	-
$F_{arch}(7, 225)$	1.28	4.04 **	0.62	-
$F_{het}(24, 214)$	1.65 *	1.33	1.77 *	-
$\chi^2_{nor}(2)$	45.61 **	32.27 **	6.11 *	-
$F_{ar}(63, 645)$	-	-	-	1.43 *
$F_{het}(144, 1229)$	-	-	-	1.30 *
$\chi^2_{nor}(6)$	-	-	-	86.77 **

Panel C: Likelihood ratio tests results Hypothesis	χ^2 statistics	Degrees of freedom
H_1: $\beta_{11} = 0$	28.664 **	1
H_2: $\beta_{21} = 0$	20.262 **	1
H_3: $\beta_{31} = 0$	7.1524 **	1
H_4: $\beta_{41} = 0$	11.944 **	1
H_5: $\beta_{11} = 0$	24.229 **	1
H_6: $\alpha_{21} = 0$	3.3584	1
H_7: $\alpha_{31} = 0$	1.3812	1
H_8: $\beta_{11} = 1$; $\alpha_{21} = \alpha_{31} = 0$;	4.9017	2

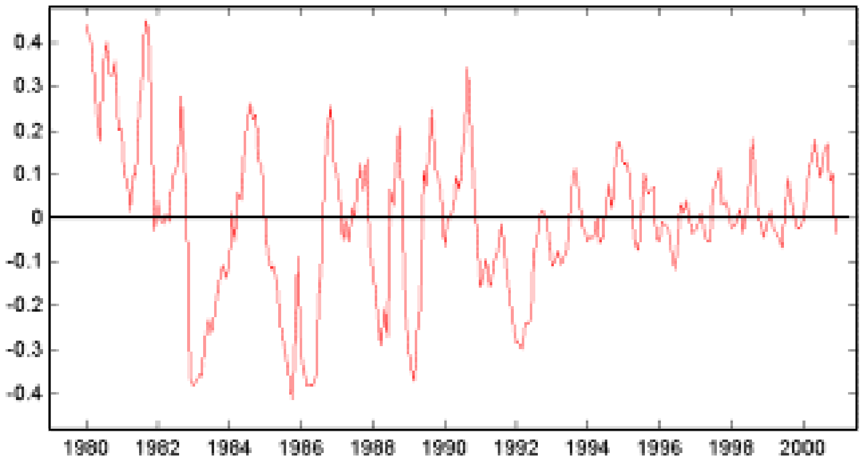
Panel D: Half-lives and AR(1) coefficients	*Pd – Pr*	*Pd – Pi*	*Pr – Pi*
Half-life	6.83	34.48	20.47
AR(1)	0.903	0.980	0.967

[1] Critical values from Osterwald-Lenum (1992). *(**) denotes significance at the 5% (1%) level.

Half-life indicators in panel D of Table 4 suggest that the case of beef appears to be the most interesting of all. In general, estimated half-lives are larger for beef than for the cereals. Results imply that the domestic and the regional markets are not integrated to the international (AU-NZ) market, as estimated half-lives imply a very low speed of adjustment. This confirms the results of the cointegration analysis where the regional price appeared to be an independent force, while the international (AU-NZ) price appeared with the wrong sign in the cointegrating vector. However, the estimated half-life between the regional price and the domestic price (less than seven months) implies an elevated grade of market integration between these markets.

From the restricted cointegrating vectors we can analyze the condition of long-run equilibrium relationships. For reasons of space, only the case of beef is reproduced in Figure 1. The estimated cointegrating vector exhibits a reduction in the fluctuation range after 1991 that can be associated with the liberalization of cattle exports in Uruguay, potentially generating a stronger link between the domestic price and the regional price.

Figure 1. Estimated cointegrating vector: Beef



V. Concluding remarks

The results of the cointegration analysis identify different patterns of long-run market integration in the commodity markets considered. Specifically, in the case of sorghum and maize, the three markets considered are found to be spatially integrated. In the case of sorghum, the US Gulf Ports price is the primary source of information, while in the case of maize, the Argentine and US Gulf Ports prices share price leadership. In the case of wheat, evidence of perfect market integration between the regional and international markets was found, however, no evidence was found of spatial market integration between the Uruguayan market and the regional or international markets. Finally, in the case of beef, there is no connection between the Argentine and the AU-NZ markets, while there is some evidence of market integration between the Uruguayan market and both regional and international markets.

From the half-life persistence indicators analysis, further evidence on the pattern of market integration was found. Specifically, price signals from international markets are transmitted at a faster speed to the regional markets than to the domestic markets, implying a larger degree of market integration of Argentine markets than Uruguayan markets. In the case of the domestic markets, price signals are transmitted at faster speed from the regional markets than from the international ones, confirming that the Uruguayan economy is highly dependent on the region. Consistent with this analysis, it was also found that there is no great difference between the wheat market and the other cereals markets, as expected. In the case of beef, results suggest that the Uruguayan and Argentine markets are not integrated to the AU-NZ market, as estimated half-lives imply a very low speed of adjustment. However, the domestic price appears to be strongly connected to the regional price.

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