COMMERCIAL LIBERALIZATION IN BRAZIL AND INTEGRATION IN THE MARKETS OF AGRICULTURAL COMMODITIES: THE COTTON, CORN AND RICE MARKETS

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ABSTRACT

The objective of this work is to analyze the long-run relationship between series of prices in some of the main export markets and Brazilian prices for cotton, rice and corn. Our aim is to verify if reductions in trade barriers favored by changes in the internal / external commercial relationships, as well as in the internal commercial and sectorial politics, were adequate to integrate the considered markets, as established by the proposition of the Law of One Price. The technique of cointegration time series developed by Johansen (1988) was used to decompose the long-term relationships in terms of tests on the alpha parameters (adjustment coefficients) and beta (importance of each variable in the long term adjustment process). The obtained results were not uniform in the sense of offering support to the proposition of marketing integration. For the rice and corn models, one cointegration vector was obtained. Relations of cointegration could be evidenced only between Argentina / Rotterdam for the corn model, and Bangkok / Uruguay and Uruguay / Brazil for the rice model. As for cotton no integration relationship was evidenced.

Key words: Markets integration, Law of One Price, Cointegration.

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1. Introduction

In the nineties, the globalization and economic integration events, and the internal market reforms, with the objective of promoting foreign trade (mainly by reducing tariff barriers), as well as the reduction in the extension of agricultural politics (due to the exhaustion of the public expenses model virtually in use for the sector until the mid-eighties) have been propitiating important qualitative changes in the economic environment in which the agents operate, involving reductions in the costs of trade, as well as in arbitration costs between countries.

At present, it is argued that the increase in external competition would have contributed to widespread a new crisis in some segments of the agricultural sector[3]. This fact has caused a flagrant connection between the evolution of prices of domestic consumed products and the fluctuations of international prices, as a consequence of the combination of changes in both the international and national economic scenarios (due to the commercial opening and marked appreciation in the Brazilian currency).

The present work is intended to find out to what extent the internal and external prices of products directed to the internal market and products of which Brazil operates in the international market as an importer, have presented a long period of common behavior, once, in along this decade, Brazil has been as an important net importer from countries of Mercosur as well as from the world market.

Therefore the empirical analysis will follow the cases of the corn, rice and cotton. We intend to infer specifically if the existence of long run relationships between the internal and external prices is enough for the markets to be characterized as being spatially integrated. In this context, integrated markets are usually those in which prices are determined in an interdependent way (Faminon and Benson, 1990) in the sense that price changes in a market are transmitted to the prices in other markets.

In international relationships, the proposition of the Law of One

[3]These worries represented the chief themes in the debate occurred in the National Seminar on Rural Revenue and Employed in Agriculture (Commission of Agriculture – Chamber of Representatives, September, 1997).

Price (LOP) in the concept of integrated markets, which in the absence of barriers to trading the arbitration mechanisms will guarantee that the price of a very homogeneous commodity in different countries, expressed in common currency, cannot differ in a value that is higher than the transaction costs. That relationship can be interpreted as a relationship that becomes true in the long term, without excluding the possibility that deviations may happen in the short term.

2. The world markets of rice, corn, cotton and the Brazilian imports

In the world corn market, the United States stands out as the largest producer. Brazil, however, occupies the third place in world production, with an output of around 31 million tons, but has been falling back in its production, having to import systematically to complete the internal provisioning, with an imported quantity representing around 5% of the consumption needs in 1998. Among the main corn suppliers to Brazil, Argentina stands out as the largest one, with a participation of around 41% on the total of Brazilian imports; 41% from 1994 to 1997 and 89% in 1998 (on average)[4].

The participation of other exporters, as the United States and the European Union is decreasing, whereas Paraguay, which did not export to Brazil until the beginning of the decade, participated with about 9% of the Brazilian imports, on average, from 1994 to 1997.

In the world rice market, Brazil stands out as a net importer, holding the third position in the ranking of the largest importers. The participation of imports for domestic consumption has been of around 14% in the 1997/98 crop and of 11% in the 1998/99 crop. Up to 1990, when the Brazilian market was protected from foreign competition by high import tariffs, Brazilian imports were basically from Asian countries and from the United States. With the consolidation of Mercosur and the opening of the Brazilian

[4]This data are from National Company of Provisioning (CONAB).

economy to the external trade, Argentina and Uruguay increased their participation in the domestic Brazilian market from 52% in 1991 to 91,7% in 1997.

The world production of cotton, around 22,2 million tons, has the United States, India and China as the largest producers. In Mercosur, Argentina is the largest producer and one of the ten largest ones in international market in the commodity. The collapse of the Eastern Europe's and the old Soviet Union's economies are factors that imprinted a strong tendency on falling prices. In the domestic market, the productive segment of this crop was one of the most affected ones by world commercial opening, since batting cotton tariffs were brought down to zero in 1990. Additionally, the credit conditions on practice for imports in foreign market contributed to its strong internationalizing trait. In this context, the proportion of Brazil's import volume for domestic consumption grew from 9,7% in 1988 to about 60% in the 1992/93 crop and also in 1996/97.

3. Data

Monthly data series were used, covering the period from January 1990 to June of 1998. The data series were chosen by taking into account some of the largest export markets in each commodity and the chief Brazilian importers from Mercosur.

The price series selected were as follow: on corn: Argentina (FOB Buenos Aires), Paraguay (FOB Assuncion), United States (Chicago - first delivery future price), Rotterdam (CIF origin in United States), and Brazil (FOB Parana). On rice: Brazil (Parana whole sale), Argentina (FOB Buenos Aires), Uruguay (FOB Montevideo) and Thailand (FOB Bangkok);· on cotton: A and B Indexes quoted in the Liverpool market, Brazil (FOB Santos) and New York (first delivery future)[5].

The data sources are the National Provisioning Company

[5] Index A is the average price of the 5 cheapest origins out of the 14 best ones in terms of quality, while Index B is the average of the 3 cheapest origin out of 8 different origins considered as the worst ones in terms of quality.

(CONAB) for the corn prices series in Brazil and Chicago, cotton series in New York, rice series in Bangkok and cotton series in Brazil. Other sources were from the Ministry of the Agriculture, Livestock and Fishery of Paraguay (corn series Paraguay), Ministry of the Agriculture, Livestock and Fishery of Uruguay (rice series - Uruguay), the National Institute of Statistics and Censuses of Argentina (INDEC) for Argentina series of (rice and corn) and the Department of Agriculture of United States (USDA) for the Rotterdam corn series. A and B indexes for cotton's international prices are present in cotton outlook of Liverpool. All the price series were expressed in reais (Brazilian currency), using for its conversion the nominal currency exchange rate (real-dollar).

4. The analytic model

The L.O.P. expresses a relationship of long-term balance among the prices established in two or more different markets. For a price change in the price in one market there is a change in the balance price on the other related markets. The basic relationship used expresses L.O.P. for two different countries as:

$$P_{it} = a + bP_{jt} + u_t \qquad (1)$$

where P_{it} and P_{jt} are the prices of a certain commodity in the markets of two countries i and j, for a period of time t.

In cases where the temporary series are not stationary (because they have a unit root) when considered individually, but a linear combination between the same ones being stationary [or integrated of order zero– expressed by I(0)], the cointegration identification is obtained among the variables, in a way as to get a stable relationship in long term, Engle and Granger (1987)[6]. Tests of unit roots for the individual series involved in each model should be completed before cointegration tests

[6] For detailed approaches on this methodology see Hamilton (1994) or Harris (1995).

presented as follows.

4.1. Determination of the integration order - unit root test

Stationary series present both, average and variance constants through time. The number of times a series should be differentiated ($\Delta y_t = y_t - y_{t-1}$) to become stationary can be indicated by the integration order of a variable (or for the number of unity roots). In this work the procedure used was the standard unit -root tests (Dickey Fuller) based on the following regression:

$$\Delta y_t = \alpha + \beta t + \gamma y_{t-1} + \sum_{i=1}^{k-1} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad (2)$$

$$\gamma = \sum_{i=1}^{k} \rho_i - 1 \qquad \text{and} \qquad \gamma_i = -\sum_{j=i+1}^{k} \rho_j$$

where k is the order of the auto-regressive process that describes the behavior of the temporary series and it should be sufficiently high to assure that the residuals are not correlated, r is the root associated to the variable dependent. Y denotes the dependent variable and Δ denotes the difference operator. For the choice of a k value a general specification was assumed, with 12 lags on the dependent variable, and successive adjustments being taken to include the lag numbers that result in smaller values for the Akaike's approach, according to Luthekpohl (1991).

The parameters to be estimate are the α, β, γ *e* γ_i. The statistics τ_τ τ_μ and τ, presented by Dickey & Fuller (1981) correspond to test t for the estimate of the y_{t-1} variable coefficient in equation (2). Those statistics are specified for a model such as (2), but they differ in the deterministic components they contain. This way, τ_τ tests the significance of that variable for the model that includes a constant and a tendency, just as represented in (2) and from here forward designated as model I. Thus, the statistics Φ_3 and Φ_2 provide verification respectively for the joint hypothesis: i) existence of unit root without tendency but with a possible term of intercept

(Φ_3), and ii) existence of unit root and a term of intercept. Statistical $\tau_{\beta\tau}$ tests only the significance of the tendency term.

Statistical τ_μ tests the significance of y_{t-1} in the model only with constant (here denominated model II); in the same model, the Φ_1 test the hypothesis of the presence of unit root and non existence of an intercept and $\tau_{\alpha\mu}$ only tests the significance for the term of the intercept; also statistical τ tests the presence of a unit root in the model without deterministic components (here designated as model III). The hypotheses tested in all three models correspond to a null hypothesis that the series is not stationary [H_0: Yt is not I(0)]; against the alternative hypothesis that the series is not integrated, or else, it is a stationary series [H_1: Yt is I(0)].

4.2. The Johansen procedure for cointegration

Suppose the vector of p variables (*p x 1*), represented as the expression $Z = (Z_{1t},\ldots,Z_{pt})$ in price variables in the approached countries - being considered that this vector assumes the following auto-regressive process (VAR) of K order which is written as an Vector Error Correction Model expressed through a matrix - adopts the following form:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \cdots + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi Z_{t-k} + \Psi D_t + \mu_z + \varepsilon_t \quad (3)$$

Where $\mathbf{D}_t$ are dummy variables that take into account shocks of short run and seasonal dummies, recommended when the data present smaller periodicity than the annual; μ_t represents the constant term and ε_t is a vectorial *p x 1* of random errors distributed identically and independently.

This form of specification of the system contains information on the adjustments of short and long run for changes in *Zt*, through estimates of Γ_i and Π respectively (Harris, 1995, p.77). Matrix Π is of order *p x p* and it embeds information about long run relationships among the variables. It is formed by the product of matrix α and β, [$\Pi = \alpha \beta'$], being α the adjustment speed to a given unbalance (also called matrix of coefficients adjustment) while β is a matrix of long run coefficients or cointegration matrix. Both matrices α and β have dimensions *p x r*, where

r is the number of cointegration relationship.

If the linear combinations expressed by $\boldsymbol{\beta' Z_t}$ are stationary, the rank of Π is given by $r \leq p$, where r determines the number of distinct cointegrating vectors that can exist among variable p included in the system. The nullity hypothesis is expressed by: H_0: **rank** $(\Pi) \leq \mathbf{r}$ or $\Pi = \alpha\boldsymbol{\beta}'$. The Trace test is applied to verify the existence of the maximum number (r) of cointegrating vectors, and the test of Maximum Eigenvalue shows the existence of exactly r vectors of cointegration against the alternative of existence of $r+1$ vectors. Both are defined by:

$$\text{The Trace Test} = -T\sum_{i=r+1}^{p}\ln(1-\hat{\lambda}_i) \quad \text{with } r = 0,1,2,\ldots, p-1 \qquad (4)$$

$$\text{The Maximum Eigenvalue Test} = -T\ln(1-\hat{\lambda}_{r+1}) \qquad (5)$$

Being T the number of observations and $\hat{\lambda}_i$ the estimated eigenvalues.The critical values for the statistical tests used in this work are presented in Osterwald-Lenum (1992).

4.3. Tests of hypotheses on the parameters α and β

Tests of hypotheses on the parameters beta and alpha allow to formulate and to test more consistent hypotheses in the economic point of view. They are presented and discussed in Johansen and Juselius (1990 and 1992) and they have not been much applied in recent researches. Among the works that applied them are Larue and Babula (1994), Lopes (1996), and In and Inder (1997).

Matrix β presented above contains the p coefficients in each cointegration relationship. β_i characterizes the relationship that must be assured between the variables when in a long-run equilibrium. They can be regarded as the relative amounts by which the level of each cointegrated variable must converge as the system re-balances or "error-corrects" itself after a shock, towards the long-run-balance pattern (Larue and Babula, 1994).

The hypotheses on parameters beta are as follows (Johansen and Juselius, 1990): H_1: $\beta = \mathbf{H}\varphi$, where **H** represents a matrix of dimensions (p x s) and s represents the number of coefficients β that are not restricted [alternatively matrix $\mathbf{H}^*$ is defined when the cointegration model contains a constant term (Johansen and Juselius, 1990, p.172)]. Matrix φ (s x r), with parameters that should be estimated with **r** cointegration vectors. The likelihood test reason is given by the expression:

$$\xi_r = T \sum \ln[(1-\lambda^*_i)/(1-\lambda_i)] \qquad \text{with } i = 1,\ldots, r \qquad (6)$$

The presence of asterisks (non-asterisks) generates models with (without) the restriction imposition in the tests. In this case, the analysis involves a space I(0), conditioned to a number of cointegration relationships (r) previously selected.

This research tests two key hypothesis associated to those parameters β: *i*) it tests for the relevance of each variable in the cointegration space (exclusion of the variable of the cointegration space) and, *ii*) among the variables whose tests on the previous hypothesis pointed to the permanence in the cointegration space, it should be tested if the integration degree among the markets is high enough as to conclude that the markets are perfectly integrated and, therefore, that the Law of One Price is effective.

The perfect integration of markets can be expressed in two different ways. The first one consists of defining as perfectly integrated markets those where a representative price exists for a group of markets as a consequence of the arbitration actions on prices in the international market. In this sense, the existence of a representative price for a group of markets is associated to the presence of multiple cointegration vectors, according to Goodwin (1992) and Lopes (1996). Rivera and Helfand (1999) note that the existence of trade flows connecting n places, and the presence of cointegration relationships, specifically with n-1 cointegration relationships, supply enough conditions to say that the involved markets are characterized as integrated markets.

A second interpretation consists of defining perfectly integrated markets as those where a variation in the price of one market is transmitted to the rest of the markets in a complete way. If in a context of multiple variables, for example, as in the case of the model for rice, variations are verified in the prices of the Argentinean market they are completely transmitted and, in the same proportion to Brazilian market, it is assumed that the balance in long term can be represented by the expression:

$$\beta' Z_t = \beta_{Ban} P_{Ban,t} + \beta_{Br} P_{Br,t} + \beta_{Uru} P_{Uru,t} + \beta_{Arg} P_{Arg,t} \qquad (7)$$

where the subscript associated to the coefficients of long run and prices, Ban, Br, Uru and Arg refer respectively to the coefficients and prices of the series of Bangkok, Brazil, Uruguay and Argentina. For the variations in prices of Buenos Aires's market, in Argentina to be completely transmitted and in the same proportion as the Brazilian market, it is assumed that the elasticity of the second in relation to the first should be as one. Taking in consideration that the variables are in logarithms and that in the equilibrium $\beta'Z=0$ the result is:

$$\frac{\partial P_{Brasil}}{\partial P_{Argentina}} \; \frac{P_{Argentina}}{P_{Brasil}} = \frac{-\beta_{Argentina}}{\beta_{Brasil}} = 1$$

Therefore, the perfect integration of the markets will be verified if $\beta_{Argentina} = -\beta_{Brazil}$. In this case, the cointegration vector can be represented in the following way (*, -1, *, 1, *), where the parameters are exposed obeying the following disposition: Bangkok, Brazil, Uruguay, Argentina and intercept term. The asterisks indicate that the corresponding parameters β_i are not restricted. In this context, the perfect integration among equal markets is tested starting from the following general hypothesis:

H_0: $\beta_i = -\beta_j$ $(i \neq j)$

Matrix **H*** for the particular case of the perfect integration test among the Argentina and Brazil markets is expressed by (8). An alternative form of representing this hypothesis consists of formulating the cointegration vector in the way (0,-1, 0, 1, *), in which the other cointegration parameters were equaled to zero (the asterisk represents the constant term). In this case, the test points to existing relationship in Argentina and Brazil market prices forming a stationary relationship in themselves and forming a separated system in which the other prices do not intervene. This test is made by means of the definition of main **H*** in (9):

$$H^* = \begin{vmatrix} 0 & 1 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{vmatrix} \quad (8) \qquad H^* = \begin{vmatrix} 0 & 0 \\ -1 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 1 \end{vmatrix} \quad (9)$$

The hypotheses concerning the alpha parameters (or coefficient of adjustment speed) can be represented in two ways:

$$\alpha = A\phi \qquad \text{or} \qquad (B'\alpha = 0) \qquad (10)$$

where it is **A** is a matrix $(p \times m)$ and m is the number of non- restricted coefficients; ϕ is a matrix $(m \times r)$. This same hypothesis can be expressed in the exposed way in parentheses for specifying a matrix $\boldsymbol{B}$ of dimensions $[p \times (p - m)]$ such that $\mathbf{B'A} = \mathbf{0}$ and $(p\text{-}m)$ equals the number of lines of restrictions imposed on α. The likelihood function for the test to this hypothesis can also be represented by expression (6) where the estimated eigenvalues are generated with (without) the restrictions imposed to parameters α.

Parameters alpha provide two types of information depending on its significance and magnitude. The significance indicates that the variable price (to which the parameter is associated), is not a weak exogenous in relation to long-run parameters β. The weak exogenous variable means that the same does not react according to changes in the long term relationship equilibrium. The magnitude of parameter α indicates the adjustment speed of the respective variable price associated to it, directed to the long run equilibrium. A small value of alpha indicates that, facing a situation of transitory unbalance, the respective variable price is adjusted slowly to return to the pattern of long run balance. A high coefficient, on the contrary, indicates that this it is quickly produced (Harris, 1995). The individual significance of the parameters alpha, is tested starting from the formulation of the null hypothesis: H_0:α_i=0;being matrix **B'** for the particular case of the test H_0: $\alpha_{Brazil} = 0$ in the rice model: **B' = [0 1 0 0]**.

A second possible hypothesis consists of verifying the equality of parameters α_i, or, the ones equivalent, the equality in the answer speed of the different variables, in a situation of short run unbalance of in the adjustment process to the long run pattern. The null hypothesis is as: H_0: $\alpha_i = \alpha_j$ (i # j) . Matrix **B'**, for example, adopts the following form: B' = [0 1 0 -1] to test the hypothesis H_0: $\alpha_{Brazil} = \alpha_{Argentina}$.

5. Results and discussion

5.1. Tests of unit root

The tests of unit root for the corn and rice series will be found in Table 1, and for the cotton series in Table 2. In these tables, models (I), (II) and (III) are those presented in the methodology.

For the corn series, the results enable us to verify that the series of Paraguay prices are stationary and will be excluded from the cointegration relationship, while all the others are not stationary and integrated of order one.

In the case of the series of prices for rice, although the statistical τ_τ and τ_μ (at 5% of significance) and ϕ_2 indicate the possible stationarity of the Brazil series, the results of statistics ϕ_3 and ϕ_1 for this series do not let us reject the joint hypothesis of the existence of unitary root without tendency and a possible intercept term, so that it cannot be affirmed in an unequivocal way that this series is stationary. In the context, it is considered that all series, including Brazil series, are not stationary and integrated of order one

The results of the unity root tests for cotton prices series reveal that the New York and Index A series are stationary. In the context, Index B and Brazil series will be considered in the cointegrating analysis.

Table 1 - Results of the tests for unitary root for the corn and rice series 1990 /01 - 1998/06

MODEL	STATISTICAL	Critical values		SERIES OF CORN PRICES[1/]				
		(5%)	(1%)	Argentina (1)	Chicago (4)	Rotterdam (1)	Paraguay (2)	Brazil (7)
I	τ_τ	-3.45	-4.04	-3.44	-2.75	-2.96	5.00*	-3.38
	ϕ_3	6.49	8.73	4.17	2.61	3.15	8.97**	4.13
	$\tau_{\beta\tau}$	2.79	3.53	-2.89*	-1.72	-2.58	4.49**	-3.13*
	ϕ_2	4.88	6.50	6.20*	3.79	4.64	13.44**	5.73*
II	τ_μ	-2.59	-3.51	-1.93	-2.11	-1.57	-	-1.22
	$\tau_{\alpha\mu}$	2.54	3.22	5.15**	2.06	1.54	-	1.13
	ϕ_1	4.71	6.70	1.91	2.38	1.32	-	1.17
III	τ	-1.95	-2.60	1.55	0.68	-0.52	-	-1.02
II nas $\Delta\Delta Y_t$	τ_μ	-3.45	-4.04	-7.16**	-	-	-	- 4.69**
III $\Delta\Delta Y_t$	τ	-1.95	-2.60	-	-7.16**	6.45**	-	

1 / the number in parentheses refers to the order of lags order used in the regression

The * indicates that the null hypothesis is rejected at 5%; * * they indicate rejection at 1%; $\Delta\Delta Yt$ represents the suitable model obtained through second differences. The critical values are found in Dickey and Fuller (1981, p. 1062-1063).

Tabela 1 - continuation

MODEL	STATISTICAL	Critical values		SERIES OF RICE PRICES[1/]			
		(5%)	(1%)	Bang-kok (3)	Brazil (1)	Uruguay (1)	Argentina (2)
I	τ_τ	-3.45	-4.04	-3.35	-3.83*	-2.05	-3.56*
	ϕ_3	6.49	8.73	3.75	5.08	1.48	4.41
	$\tau_{\beta\tau}$	2.79	3.53	-2.26	-2.69	-1.62	-3.34*
	ϕ_2	4.88	6.50	5.61*	7.56**	2.13	6.60**
II	τ_μ	-2.59	-3.51	-2.41	-2.71*	-1.26	-1.34
	$\tau_{\alpha\mu}$	2.54	3.22	2.40	2.67*	1.22	1.32
	ϕ_1	4.71	6.70	2.93	3.74	0.88	0.91
III	τ	-1.95	-2.60	1.20	1.09	-0.53	-0.24
II nas $\Delta\Delta Y_t$	τ_μ	-3.45	-4.04	-	- 9.98**	_	-6.19**
III $\Delta\Delta Y_t$	τ	-1.95	-2.60	-4.51**	-	7.78**	-6.21**

1 / the number in parentheses refers to the order of lags order used in the regression
The * indicates that the null hypothesis is rejected at 5%; * * they indicate rejection at 1%; $\Delta\Delta$Yt represents the suitable model obtained through second differences. The critical values are found in Dickey and Fuller (1981, p. 1062-1063).

Table 2 - Tests of unitary root (Increased Dickey-Fuller) for cotton series,1990/01 - 1998/06.

MODEL	ESTATÍSTC	Critical values (5%)	Critical values (1%)	A[1/] Index (4)	B[1/] Index (9)	New York[1/] (2)	Brazil[1/] Santos (2)
I	τ_τ	-3.45	-4.04	-4.08**	-2.63	-4.41**	-2.45
	ϕ_3	6.49	8.73	5.85	2.98	6.49*	2.01
	$\tau_{\beta\tau}$	2.79	3.53	-3.73**	-2.75	-3.58**	0.43
	ϕ_2	4.88	6.50	8.39**	3.88	9.74**	3.01
II	τ_μ	-2.59	-3.51	_	-0.41	_	-2.42
	$\tau_{\alpha\mu}$	2.54	3.22	_	0.36	_	2.43
	ϕ_1	4.71	6.70	_	0.63	_	2.96
III	τ	-1.95	-2.60	_	-1.07	_	-2.12
III series nas $\Delta\Delta Y_t$	τ	-3.45	-4.04	-	-5.52**	-	-8.07**

1Value in parentheses indicates the number of lags used in the regression. * indicates that the null hypothesis is rejected at 5%, * * indicates rejection at 1%. $\Delta\Delta Y_t$ represents the suitable model obtained through second differences. The critical values meet in Dickey and Fuller (1981, p. 1062-1063).

5.2. Cointegration tests

The empirical analysis start from the formulation exposed in (3), with four equations for the corn case, 4 equations for the rice case, 2 equations for the cotton series.

The election of the lags numbers (k) of the model took into account the approach of Akaike for multivariate models. The results indicated that, for corn, the approach of Akaike is minimized with 3 lags, for cotton, with 2, and for rice, with 12 lags. Eleven centered variables dummies were also placed in the model for rice and cotton, developed to take in consideration the stationarity presence in the short run space[7].

In Table 3, the results of the statistical λ-Max and test line are presented. The results reveal that the system formed by the corn and rice prices contain an only cointegrating vector that is equivalent to saying that in each system of prices a column of β forms an independent and stationary lineal combination of variables in **Zt,** with 3 non-stationary vectors $(p - r)$ in the case of corn and 3 in the case of rice. In the model formulated for the cotton series, the results presented in Table 4 show that there is no cointegrating relationship among the two considered series, which are Index B and Brazil.

On the number of cointegration relationship found, for Dickey, Jansen and Thornton(1994), cointegrating vectors can be understood as representing constraint on economic system which imposes a barrier on the movement of the variables in the system in the long run. Thus, more cointegrating vectors make the system more stable being, therefore, desirable that an economic system is stable in as many directions as possible. On the other hand, when only a cointegrating vector is had in a system with variable p, the system can float around p-1 to independent directions but only one is stable.

[7] This procedure was also used by Johansen and Juselius (1992) to be given on a quarterly basis, In and Inder (1997) and Lopes (1996) for monthly data.

Table 3 - Cointegration tests results for the corn models and rice, 1990/98

	Corn series					Rice series				
H_0: r	Eigen-value	λ-Max	Critical V. (95%)	Trace test	Critical V. (95%)	Eigen-value	λ-Max	Critical V (95%)	Trace test	Critical V. (95%)
0	0.40	50.33*	28.14	77.84*	53.12	0.32	34.60*	28.14	63.31*	53.12
1	0.15	15.65	22.00	27.52	34.91	0.18	17.65	22.00	28.71	34.91
2	0.10	10.80	15.67	11.86	19.96	0.08	8.08	15.67	11.06	19.96
3	0.01	1.07	9.24	1.07	9.24	0.03	2.98	9.24	2.98	9.24

The critical values are taken from Osterwald-Lenum (1992)

Table 4 - Cointegration tests results for cotton series (1990/98)

H_0: r	Eigenvalue	λ-Max	Crítical value (95%)	Trace test	Crítical value (95%)
0	0.022	2.20	15.67	2.41	19.96
1	0.002	0.21	9.24	0.21	9.24

The critical values are taken from Osterwald-Lenum (1992)

5.3. Tests of hypotheses on the alpha and beta parameters

Analyzing, initially, the significance of beta parameters for cointegrating vectors, that is to say, the hypotheses that the $\beta i = 0$, the rice model presents the indication of non-rejection to this hypothesis for the Argentina series (Table 5) only. In the corn model, the same result is evidenced for the Brazil and Chicago series. This indicates that the changes of prices in Brazil and in the future market of Chicago in the case of corn, as well as the levels of prices in the Argentinean rice market, they are not significantly important in the establishment of the pattern of balance in the long run among the other markets considered in the analysis. It also indicates that prices impacts in these markets are not capable of promoting significant adjustments in other prices considered in the analysis. Therefore, such variables can be excluded of the long run relationship.

As follows, the hypothesis of perfect integration is analyzed among the markets for which the statistic tests lead to the rejection of the nullity hypothesis for the respective beta parameters. Without imposing any restriction to other not tested beta parameters, the results do not allow us to reject the null hypothesis of perfect integration between the Argentina and Rotterdam markets in the corn model, as well as between the Thai markets (Bangkok) and Uruguayan and between the latter and the Brazilian markets, in the rice model. That is to say, these two markets can be considered perfectly integrated, in the sense that alteration of prices in a market is completely turned over to the other.

In the case of corn model, the evidence of perfect integration between the market exporters from Argentina and Rotterdam is not altered when zero restriction is placed on the other not tested beta parameters, that is, for these markets, the Law of the One Price is perfectly verified. For the model involving the rice series, on the other hand, when zero restrictions are introduced to the other untested beta parameters, the results lead us to the rejection of the hypothesis of perfect integration among the Bangkok/ Uruguay and Uruguay/Brazil markets (Table 5). It can be said in this case, that such markets do not form a stationary relationship to each other independently of another markets, that is to say, these sets of markets are perfectly tuned in a context in which the other prices intervene.

Table 5 - Tests on the significance of the β parameters and perfect integration among sets of markets. Corn and rice series, 1990/98.

Model for corn series			Model for rice series		
H_0	Likelihood ratio	Critical Value χ^2 (5%)	H_0	Likelihood ratio	Critical Value χ^2 (5%)
$\beta_{Brazil} = 0$	0.07	3.84	$\beta_{Brazil} = 0$	9.09*	3.84
$\beta_{Chicago} = 0$	1.07	3.84	$\beta_{Bangkok} = 0$	16.28*	3.84
$\beta_{Rotterdan} = 0$	5.34*	3.84	$\beta_{Uruguay.} = 0$	6.84*	3.84
$\beta_{Argen} = 0$	22.13*	3.84	$\beta_{Argentina} = 0$	0.76	3.84
$\beta_{Argent.} = -\beta_{Rotterd}$ 1/	0.55	3.84	$\beta_{Bangkok} = -\beta_{Brasil}$ 1/	12.77*	3.84
$\beta_{Argent.} = -\beta_{Rotterd.}$ 2/	2.08	5.99	$\beta_{Uruguay} = -\beta_{Brasil}$ 1/	0.17	3.84
-	-	-	$\beta_{Uruguay} = -\beta_{Brasil}$ 2/	25.68*	7.81
-	-	-	$\beta_{Bangkok} = -\beta_{Uruguai}$ 1/	0.05	3.84
-	-	-	$\beta_{Bangkok} = -\beta_{Uruguai}$ 2/	24.73*	7.81

1 / without restrictions in the not tested beta parameters

2 / restricting the other parameters b to zero

Table 6 – Tests on the significance of parameters α(exogeneity tests)
I model for the series of corn 1990/98.

H_0	Without restriction on Parameters β		With restriction $\beta_{Argentina} = -\beta_{Rotterdam}$		Otherrestrictions	
	Likelihood ratio	Critical Values χ^2 (5%)	Likelihood ratio	Critical Values χ^2 (5%)	Likelihood ratio	Critical Values χ^2 (5%)
$\alpha_{Argen} = 0$	6.17^*	3.84	6.39^*	5.99	-	-
$\alpha_{Brazil} = 0$	18.60^*	3.84	19.46^*	5.99	$18.91^{*\,2/}$	5.99
$\alpha_{Chicago} = 0$	1.48	3.84	1.76	5.99	$2.01^{3/}$	5.99
$\alpha_{Rotterdam} = 0$	2.34	3.84	3.10	5.99	-	-
$\alpha_{Brazil} = \alpha_{Argentina}$	15.06^*	5.99	3.74	5.99	$15.10^{2/}$	5.99
$\alpha_{Rotterdam} = \alpha_{Argentina}$	0.71	5.99	1.14	5.99	-	-

2 / conditioned to βBrazil = 0
3 / conditioned to βChicago = 0

The results for the alpha parameters tests are discussed as follows. In the pattern model of balance in corn prices, Chicago and Rotterdam variables are revealed weak exogenous (Table 6). In the case of the first series, the non significance of the beta parameter previously verified had already revealed that the same non significance, does not participate in the cointegration relationships, in such away, that it will normally not react to any disturbance.

In the case of the non significance of parameter alpha for the Rotterdam series, such a result is not expected, bearing in mind that this being market is an important exporter and participate in the cointegration relationship. A possible interpretation for such a result is that, although the level of this price influences the levels of prices in other markets (it could increase the levels of prices of these markets) it is not, however, significantly affected by them and is not adjusted before transitory disturbances. The objective of test $\alpha_{Chi.} = \alpha_{Rot}.= 0$ is to verify if Chicago and Rotterdam they can be considered exogenous altogether, since each series is individually so. The result shows that this proposition cannot be rejected.

Since a priori the Argentina and Rotterdam export markets are considered perfectly integrated (given the non rejection of the hypothesis βArgentina = -βRotterdam), we can ask if both of them react to an upheaval at identical speeds, even having in mind that Rotterdam is exogenously weak. The likelihood test ratio shows that this hypothesis cannot be rejected.

In this same model, another surprising result refers to the quite significant value of the α test given to Brazil series, the information that this series could be excluded from the cointegrating space (βBrazil = 0), it should be understood that there is no long run relationship between the price levels in this country and in the others. The hypothesis of weak exogeneity is already clearly rejected (and with the largest absolute value for alpha indicating larger adjustment speed to transitory unbalances). That is to say, the levels of prices in Brazil, even not significantly important in the establishment of long run balance of other markets, react with relative efficiency to transitory unbalances that happen to other prices.

Table 7 - Tests on the significance of parameters α (exogeneity tests).
Model for the series of rice 1990/98.

H_0	Without restriction on the β parameters: Likelihood ratio	Critical Value χ^2 (5%)	Restriction 1/ $\beta_{Bangkok}=\beta_{Uruguay}$	Restriction 1/ $\beta_{Uruguay}-\beta_{Brazil}$ Likelihood ratio	Critical Value χ^2 (5%)
$\alpha_{Argentina}=0$	5.04*	3.84	7.40*	6.32*	5.99
$\alpha_{Brazil}=0$	1.50	3.84	2.32	2.48	5.99
$\alpha_{Bangkok}=0$	2.26	3.84	3.65	4.51	5.99
$\alpha_{Uruguay}=0$	7.13*	3.84	7.18*	7.34*	5.99

1 / the critical values are in the last column

As Brazil has been importing significant volumes of Argentina corn, it is tested if both markets return to a given upheaval at statistically equal speeds. This proposition can be rejected.

In case of the rice market involving the series approached in this work, the significance of the parameters alpha for the of Uruguay and Argentina prices series reveal that these are not weak exogeneity for the parameters of interest, while Brazil and Bangkok series present weak exogeneity with regard to the long run balance. These results do not vary when restrictions of perfect integration are introduced among the Thai and Uruguayan markets and between this last one and the Brazilian market.

In other words, it is considered that the Argentinean market, though it is not significantly important in the establishment of long run balance for the group of the considered prices, reacts significantly to the coming unbalances of the other markets. Nevertheless the exogeneity of Brazil's and Bangkok's prices mean that these prices are not affected by deviations that may occur, in the short run, in regard to a balanced situation in other markets. In the case of Brazil, it could be indicating that the internal prices give no answers to short period unbalances in foreign markets. The exogeneity of Bangkok can be associated to the fact that Thailand is the largest producer and world exporter of rice. For that reason, it marks the evolution of the other markets but does not react significantly to other market's short period upheavals.

6. Conclusion

The study represents an attempt to discuss and obtain inferences over to what extent the internal prices of rice, corn and cotton commodities have been following the long-period pattern of external markets in a decade in which important economic transformations in the external and internal markets promoted important changes in the formation dynamics of agricultural commodities prices. Were the alterations dictated by the world commercial opening and lesser interference of the State in internal markets enough to enact efficient arbitration actions? Is the Law of One Price validated and therefore considered integrated to the market? What are the implications to the internal sectorial politics?

The cointegration procedure utilized was proposed by Johansen (1988) and the results obtained in the case of cotton did not enable us notice co-integrating relationship between the prices of Brazilian market and Liverpool Index B market (the Liverpool Index A series and New York series were not considered for they revealed stationary). In conclusion, in the case of this commodity, LOP is not assured and the markets are not integrated during the period of this analysis. We cannot say, however, that the formation of internal prices is essentially a domestic issue in this case. But, in spite of the larger commercial exchange practiced on this decade, with expressive imported volumes, there is still , in the Brazilian market, a great inflexibility towards of arbitration of the internal economic agents' prices in the external markets. An example of this refers to the difficulty that Brazilian companies in less favorable patrimonial situation have to loans for the improvement of cotton imports.

The cointegration tests for corn and rice pointed to the existence of a cointegrating vector among the series of prices considered in each model and, therefore, to the existence of a certain integration degree among the markets. This result is consistent with previous analyses (Goodwin, 1992; Lopes, 1996; and Lima and Burnquist, 1997), indicating the presence of arbitration mechanisms among the markets, and, therefore, a relationship of interdependence on prices. However, in spite of this result indicating the positive performance of arbitration mechanisms among the markets, we may that the existence of only one cointegrating vector is not a sufficient condition to say that the markets are perfectly integrated and that the Law of the One Price is perfectly verified or, alternatively, that it is verified in its absolute form. This inference is based on the arguments of Goodwin (1992b), Lopes (1996), Rivera and Helfand (1999) and Dickey, Hansen and Thornton (1994). According to these last authors, larger stability in the long period relationships among the variables of the system requests a larger number of cointegrating vectors.

The results of the hypotheses tests in the case of the corn model enabled usto verify that the vector of estimated cointegration brings an implicit relationship of perfect integration between the Argentina and

Rotterdam export markets. These also indicated the exclusion of Brazil series, meaning that this market does not endogenously participate or affect other variables in the adjustment process to the long run balance. In practical terms, the results allow us to conclude that the variables associated to the internal market still have larger relative weight in the formation of the domestic corn prices. In this case, we can refute the initial proposition that the external prices would excessively affect the behavior of the prices in the internal market.

In the case of rice, the existence of prices arbitration indicated in the cointegrating relationship, and also, from the positive significance of the Brazil series in the establishment of the long run balance pattern (pointed by the tests on the beta parameters), evidences that in fact, the internal prices have a relationship with the prices of the international market, although they do not respond to the original short run deviations of the other markets. In the same model, it was also evidenced that, in Mercosur, the Uruguayan export market and the Brazilian import market can be characterized as integrated markets in the long run. Such integration is verified in a context characterized by the connections to other prices.

The main implications of agricultural politics, in the light of the obtained results, are that, for the internal markets of cotton and corn, sectorial programs of prices politics and internal provisioning can still play an important part in the balance of the markets. It should be understood that this proposition does not mean that there is ground for initiatives in prices politics and provisioning in the molds that were effective in the last decades. On the contrary, it is necessary to take the prices, agreements and effective rules in the international market into consideration, since in the case of corn, the internal prices indications were obtained due to unbalanced external markets (significance of alpha parameter).

Specifically in the case of domestic price politics for rice, the results obtained lead us to the conclusion that the formulation of these politics should take in consideration Uruguay as an extension of the national market. Furthermore, both countries can gain significantly in the medium and long run if they get to elaborate an agricultural politics coordinated to the production aspects, supporting services and technology. Such coordination of efforts can very be important both as a better way to explore the complementarities between these countries, and , once the recovery of the capacity of Brazilian self-provisioning is reached, to make a competitive insert possible and obtain better positions for both countries in the world rice exports.

Two possible limitations to this work, which can be suggested as effective complementary works, are associated to the size of the series (due to the impossibility of obtaining larger series) and to the fact that nominal exchange rates had to be used for the conversion of the prices.

7. References

DICKEY, D.A.; FULLER, W.A. Likelihood ratio statistics for auto-regressive time series with a unit root. **Econometrica**, 49: 1057-1072, 1981.

DICKEY, HANSEN, THORNTON. A primer on cointegration with an application to Money and Income. In: RAO, B.(editor). **Cointegration for applied economist**. New York: St. Martin's Press. 1994, .9-45.

ENGLE, R.F. ; GRANGER, C.W.J. Co-integration and error correction: representation, estimation and testing. **Econometrica**, 55: 271-276. 1987.

FAMINON, M.D. BENSON, B.L. Spatial market integration. **American Journal of Agricultural Economics**. Vol. 72, no. 1, p. 49-62, Feb. 1990.

GOODWIN, B.K. Multivariate co-integration tests and the law of one price: a clarification and correction. **Review of Agricultural Economics**, vol. 14, n.2, 1992b.

HAMILTON, J. **Time Series Analysis**. Princeton University Press. 1994.

799p.

HARRIS, R.I.D. **Using Cointegration analysis in Econometric Modeling.** London: Prentice Hall/Harvester Wheatsheaf, 1995. 176 p.

IN, F; INDER, B. Long-Run relationship between World Vegetable oil prices. **The Australian Journal of Agricultural and Resource Economics.** Vol. 41, no. 4, p. 455-470. Dec, 1997.

JOHANSEN, S; JUSELIUS, K. Maximum likelihood estimation and inference on cointegration - With application to demand for money. **Oxford Bulletin of Economics and Statistics,** vol. 52, n.2, p. 169-210, 1990.

JOHANSEN, S; JUSELIUS, K. Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for UK. **Journal of Econometrics**, vol. 53, p. 211-244. 1992.

LARUE, B. BABULE, R.A. Evolving dynamic relations between the money supply and food-based prices in Canada and The United Sates. **Canadian Journal of Agricultural Economics**, vol.42, p159-176. 1994.

LIMA, S.M. A .; BURQUIST, H.L. Lei do Preço Único no mercado internacional: Testes Empíricos para Exportações do Complexo Soja (Grãos e Farelo). In: XX1V CONGRESSO BRASILEIRO DE ECONOMIA E SOCIOLOGIA RURAL. Natal, 1997. . Anais... (CD-ROM). Brasília: Sociedade Brasileira de Economia e Sociologia Rural (SOBER), 1997.

LOPES, A. I .S . **Integration Espacial de los mercados de porcino europeus**. Zaragoza, 1996. 219 p. Centro Internacional de Estudos Agronômicos Mediterrâneos

LUTHEKPOHL, H. **Introduction to multiple time series**. Berlin, Springer-

Verlag, Heidelbrg, (1991).

OSTEWARD-LENUM, M. Practitioner's Corner - A note with quintiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistic. **Oxford Bulletin of Economic and Statistics,** 54 (3): 462-472, 1992.

RIVERA,G.G.; HELFAND S.M. Spatial Relationships and market integration: the case of Brazilian rice market. (1999). In: XXXVII CONGRESSO BRASILEIRO DE ECONOMIA E SOCIOLOGIA RURAL. Foz do Iguaçu, 1999. Anais... (CD-ROM). Brasília: Sociedade Brasileira de Economia e Sociologia Rural (SOBER), 1999.