THE WORLD MARKET FOR SOYBEANS: PRICE TRANSMISSION INTO BRAZIL AND EFFECTS FROM THE TIMING OF CROP AND TRADE

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RESUMO

Este trabalho investiga a transmissão de preços no mercado mundial de soja usando econometria de séries de tempo. O modelo teórico desenvolvido por Mundlack e Larson (1992) é baseado na Lei do Preço Único e supõe que os preços se equalizam ao longo de todos os mercados locais no longo prazo, permitindo-se desvios transitórios no curto prazo. O mercado internacional foi caracterizado através de três preços relevantes: Rotterdam, Argentina e Estados Unidos. O trabalho estima a elasticidade de transmissão desses preços aos preços no Brasil. Foram realizados testes de causalidade e de cointegração para verificar se há relação de longo prazo entre as variáveis. Foi também calculada a função de resposta a impulso e a decomposição da variância dos erros para avaliar a transmissão de preços internacionais aos preços brasileiros. Aplicou-se um teste de exogeneidade para verificar se as variáveis respondem a desvios de curto prazo em relação aos valores de equilíbrio. Os resultados confirmaram a validade da Lei do Preço Único no longo prazo. Em linha com vários trabalhos, este artigo mostrou que Brazil e Argentina podem ser vistos como tomadores de preços no mercado internacional, tendo em vista que a velocidade de ajuste de seus preços em resposta a choques é maior que a verificada para os preços dos Estados Unidos, que são formadores de preço. Uma conclusão interessante foi obtida pela comparação da função de resposta de impulso ao período de colheita e de comercialização da soja no Brasil, Argentina e Estados Unidos. Estas diferenças sazonais ajudam a explicar o padrão da resposta dos preços brasileiros aos choques no mercado internacional, especialmente notando-se que a resposta a choques nos Estados Unidos é oposta à resposta aos choques na Argentina, porque a colheita nos dois hemisférios ocorre em períodos diferentes. Adicionalmente, o atraso de um mês entre as colheitas brasileira e argentina pode contribuir para a explicação de um ponto de inflexão que ocorre após um mês nas funções de impulso-resposta.

PALAVRAS-CHAVE

Soja; elasticidade de transmissão de preços; econometria de séries temporais

ABSTRACT

This paper investigates the price transmission in the world market for soybeans using time series econometrics models. The theoretical model developed by Mundlack and Larson (1992) is based on the Law of the One Price, which assumes price equalization across all local markets in the long run and allows for deviations in the short run. The international market was characterized by three relevant soybean prices: Rotterdam Port, Argentina and the United States. The paper estimates the elasticity of transmission of these prices into soybean prices in Brazil. There were carried causality and cointegration tests in order to identify whether there is significant long-term relationship among these variables. There was also calculated the impulse-response function and forecast error variance decomposition to analyze the transmission of variations in the international prices over Brazilian prices. An exogeneity test was also carried out so as to check whether the variables respond to short term deviations from equilibrium values. Results validated the Law of the One Price in the long run. In line with many studies, this paper showed that Brazil and Argentina can be seen as price takers as long as the speed of their adjustment to shocks is faster than in the United States, the latter being a price maker. An interesting conclusion was reached when the pattern of the impulse response functions was compared to the timing of crop and trade in Brazil, Argentina and the United States. These seasonal differences may help explaining the pattern of the response of Brazilian prices to shocks in the international market, especially that the response from shocks in the United States is opposite to the response from shocks in Argentina because harvest in the two hemispheres occurs in different periods. In addition, the one-month lag between Brazilian and Argentine harvests may contribute to explain a turning point in the impulse-response function that occurs one month after the shock.

KEYWORDS

Soy; elasticity of price transmission; time series econometrics

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1 – INTRODUCTION

Soybean and its derivatives enter as key inputs in several segments of the agribusiness chain. According to FREITAS et al (2001 p. 2), *"this commodity is one of the most traded agricultural products in the world, probably because of the variety of forms of consumption, which range from food (human and animal) to the pharmaceutical and steel industries. This diversity is possible because the soybean processing industry produces by-products, soymeal and oil, which constitute important input for different industrial sectors"*.

The international market for soybeans has an interesting feature of a strong degree of concentration in both the supply and demand sides. According to data from USDA (2003), 79.87% of the total world production in the period between 1994/1995 and 2002/2003 took place in three countries, United States (44.85%), Brazil (21.71%) and Argentina (13.30%). These three countries were also the world's leading exporters, accounting together for 90.50% of the 44.32 millions of tons in the same period. The United States was the leading exporter with 56.85% of the total, followed by Brazil (24.10%) and Argentina (9.54%). On the demand side, the major export destiny in the same period was the European Union (EU) with 36.98% of the total, where soy meal is basically used for animal feeding. China is the second largest market for soybean imports, with 16.22%. Together, the European Union and China account for 53.20% of world imports. It is noteworthy that the importance of China in the international market started to grow only late in the 90's, when its imports jumped from 3.85 in 1998/1999 to 10.10 millions of metric tons in 1999/2000, according to OILSEEDS (2003). As the Chinese economy maintained a strong pace of growth, its imports grew further to 21.42 millions of metric tons in 2002/2003 (OILSEEDS 2003).

Being the second largest producer and exporter of soybeans, Brazil relies strongly on the soy chain as a source of foreign currency receipts in the current account of its balance of payments. According to data from CONAB (2004), in the year 2003 Brazil exported US$ 73.1 billion in goods, of which 41.9% or US$ 30.6 billion were products of the agribusiness chain. The soy chain[7] accounted for 11.1% of total Brazilian exports, with annual receipts of US$ 8.1 billion. These data also reveal that the soy chain plays a key role not only as a source of foreign currency receipts for Brazil but also as a major source of income among the several segments of the country's domestic agribusiness chain.

Given the importance of the soy chain, several studies using time series were carried to analyze the formation of soybean prices in Brazil. Results of some of these studies are described in the remaining of this section.

PINO and ROCHA (1994) evaluated the transmission of Chicago Board of Trade (CBOT) quotations to the prices of soybeans at industry and farm levels in Brazil. In general terms, these authors reached the conclusion that domestic soybean prices in Brazil are strongly influenced by the movement of CBOT prices. The latter determines the quantity supplied in Brazil and also affects prices of domestic byproducts (grain and soymeal).

MARGARIDO and SOUSA (1998) analyzed the transmission of soybean prices to prices in Brazil and in Paraná. The study found out that the changes in CBOT quotations are only partially transmitted, but without time lags, to prices at farmer level considering Brazil as a whole and Paraná. Such low elasticity of price transmission was found to be apparently related to the strategies pursued by crushers, which choose the final destiny of the soybeans taking into account the relationship between international and domestic prices. Besides, domestic consumption of soybean derivatives is relevant since soy meal is used to feeding birds

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[7] The soy complex encompasses the soybeans, bran, gross oil, refined oil and other oils.

while oil is widely consumed internally because of its price advantage against alternatives such as rape seed and sunflower. These factors, plus the cost of producing in Brazil, possibly influence the behavior of the elasticity of price transmission, making it smaller than unity.

MARGARIDO et al. (1999) broadened the scope of soy chain studies by measuring the elasticity of price transmission involving the CBOT, the prices at the Rotterdam Port and domestic prices in Brazil and Argentina. They found out that variations in Rotterdam prices are transmitted more intensely and faster to domestic prices in Brazil and Argentina than variations at CBOT. Since Rotterdam prices are spot price for imports destined to the European Union and CBOT prices are future prices that reflect supply conditions, these results suggest that demand prices play a more significant role in the formation of Brazilian and Argentine domestic prices than supply prices. Another finding of this study was that Argentine prices are more sensitive than Brazilian prices to variations in international prices, what probably reflects the specific characteristics of the two markets. Brazil has a relatively more important domestic consumption market, while Argentina sends out the bulk of its production to international markets, being more exposed to those variations.

MACHADO and MARGARIDO (2001) using a different methodology based on Granger Causality tests reached a similar conclusion that Brazilian and Argentine domestic prices are more sensitive to variations in Rotterdam than in CBOT.

2 – THEORETICAL MODEL

MUNDLACK and LARSON (1992) developed the theoretical model used in this paper. This model shows how variations in external prices transmit to domestic prices[8]. Based on the Law of One Price[9], the domestic price of soybeans can be written as a function of the international price of the commodity, the nominal exchange rate and the specific trade policy[10]. Algebraically, this model[11] can be stated as follows:

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

where:
P_{it} is the domestic price of the i product at period t;
P^*_{it} is the international price of the i product at period t;
E_t is the nominal exchange rate at period t
Multiplying both sides of (1) by $1/E_t$, we have:

$$P_{it}^{us\$} = P_{it}^{*} \tag{1a}$$

where $P_{it}^{us\$}$ refers to domestic prices of product i taken in US dollars in time t. Thus, in this paper the term $P_{it}^{us\$}$ stands for the price of the soy beans in Brazil, however in US dollars.

The model laid out by MUNDLACK e LARSON (1992) can be modified to incorporate specific elements from the international market of a given product. In the long run, the law of the one price states that domestic prices tend to equal world market prices for any given product. In other words, the elasticity of

8 In the absence of intervention, domestic prices equal international prices.
9 KRUGMAN and OBSTFELD (1997) note that although the Purchasing Power Parity is similar to the Law of One Price, there is a difference between them. While the latter apply to a single product, the first refers to the general price level reflecting the prices of all products that compose a bundle of goods taken as a reference. However, if the Law of One Price is valid for every product in this bundle, then the Law of One Price is equivalent to the Purchasing Power Parity.
10 Since this paper is not concerned with trade policies, this variable was not included in the model.
11 This model does not take into account qualitative differences between the products and the transport costs, storage, or the prices of the nontradable domestic inputs.

transmission has to be unity. In such a context, the world market price can be written in the form of a multiplication operator, since the objective of this paper is to estimate the relevant elasticities of price transmission:

$$P_{it}^{*} = \prod_{i=1}^{n}\prod_{j=1}^{n} W_i^{j}\, P_{it}^{j} \tag{1b}$$

where the subscript index i represents product, while the subscript j is for the country and W is the respective weight in the weighted geometric average that composes the international price. Taking the case of the soy international market, its international price can be decomposed as follows:

$$P^{*}_{it} = W_i^{Rot} * P_{it}^{Rot} * W_i^{Arg} * P_{it}^{Arg} * W_i^{USA} * P_{it}^{USA} \tag{1c}$$

where P_{it}^{Rot} is the price of the soy beans in Rotterdam in dollars; P_{it}^{Arg} is the price of the soy beans in Argentina and P_{it}^{USA} is the price of the soy beans in the United States. Two assumptions are necessary at this stage. First, we must assume that the sum of the three weights equals unity in equation 1c so that equality holds in respect to equation 1c. The second assumption is that the weights in the equation have identical values, given the characteristics of the international market for soybeans. Since few countries hold a sizable share of the world supply, one should expect to find a high correlation among its prices, provided the arbitrage process is functioning properly as it is described by the Law of the One Price.

An error term (u) was added to the equation to capture possible deviations from variables not included in the model. Writing equation (1a) in the logarithmic form, we obtain:

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

or alternatively, through decomposition of the international price[12]. Thus,

$$p_{it}^{US\$} = p_{it}^{Rot} + p_{it}^{Arg} + p_{it}^{USA} + u_{it} \tag{2a}$$

Where u ~ IID(μ , σ^2) and E (p * u) = 0, from equation 2 meaning that u_{it} shows no correlation with the other explicative variables of the model. The simplest model thus assumes domestic prices of a single product to be a function of its respective international price and the disturbance term. From equation 2 the following model can be estimated:

$$p_{it}^{US\$} = \alpha + \beta\, p^{*}_{it} + \varepsilon_{it} \tag{3}$$

where α is a constant (intercept) and β is assumed to be unity. The β coefficient is then the elasticity of the domestic price in US Dollar against its international price, i.e., it is an elasticity of price transmission. When its value is unity, variations in the international price are fully transmitted to domestic prices. On the other hand, when the β value is zero, variations in international prices do not lead to any response in domestic prices, the domestic economy being a completely closed one. The most common situation is

[12] Several studies reached the conclusion that the price of soy beans in Brazil is influenced by its prices in the European Union and the United States; this means that the direction of causality is uni-directional. For Argentina, it is relevant that its harvest reaches the market with a lag of one to two months as compared to the Brazilian harvest. Thus, the estimates for Argentine harvest may influence prices in Brazil, as one takes into account that the international market for soybeans is highly concentrated. The United States, Argentina and Brazil account together for roughly 80% of world soy production, causing a high correlation between prices in these three countries.

when β lies between zero and one, reflecting a specific trade policy adopted by the country or any sort of trade restriction imposed to the market.
The model shown in equation 3 can be restated to present the international price of soybeans by country:

$$p_{it}^{US\$} = \alpha + \beta_1 p_{it}^{Rot} + \beta_2 p_{it}^{Arg} + \beta_3 p_{it}^{USA} + \varepsilon_{it} \tag{3a}$$

Equation 3a has an advantage over equation 3, namely that it allows direct estimation of the respective elasticities of price transmission. In this case we assume that $\beta = \beta_1 + \beta_2 + \beta_3$. When $\beta = 1$, then the Law of the One Price holds, meaning that price variations in the international market[13] for soybeans are fully transmitted to the domestic price in Brazil, denoting unity elasticity of price transmission. Thus, in this case $\beta_1 + \beta_2 + \beta_3 = 1$. Another hypothesis to be tested is that these four markets are integrated, thus price transmission flows without restrictions between the markets involved, or, the arbitrage system equalizes the prices in all markets in the long run. Thus, one should expect that the three coefficients be identical $\beta_1 = \beta_2 = \beta_3$.

3 – OBJECTIVES

The objective of this study was to quantify the elasticity of price transmission[14] in the international market for soybeans, involving Cost Insurance and Freight (CIF)[15] prices at Rotterdam, Free on Board (FOB)[16] prices in Brazil, Free on Board (FOB) in Argentina and U.S. NO.1 Yellow[17] Cash Central Illinois (USA) for the period between October 1995 and October 2003. There were carried causality and cointegration tests in order to identify whether there is significant long-term relationship among these three variables. It was also calculated the impulse-response function[18] and forecast error variance decomposition to analyze the transmission of variations over time in the Rotterdam Port, Argentina and United States prices over Brazilian domestic prices. An exogeneity test was also carried out so as to check whether the variables respond to short term deviations from equilibrium values.

4 – DATASET AND METHODS

4.1 – Dataset

For this paper there were built four time series with monthly observations: Cost, Insurance and Freight (CIF) prices at the Rotterdam Port (*ROT*)[19], Free on Board (FOB) prices in Brazil (*BR*), FOB prices in Argentina (*ARG*) and U.S. NO.1 Yellow Cash Central Illinois (*USA*). Basic data on soybean prices were obtained from OILSEEDS (oct. 1995 / oct. 2003). Since all variables were used in logarithmic form, coefficients can be directly read as elasticities. Logarithms of variables *ROT*, *BR*, *ARG* and *USA* were named *LROT*, *LBR*, *LARG* and *LUSA* respectively.

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[13] International market stands for Rotterdam, United States and Argentina.

[14] BARROS and BURNQUIST (1987, p. 178) note that the “elasticity of price transmission refers to the relative variation in the price at a market level in relation to the variation in the price at another level, keeping equilibrium in these two levels after the initial shock in each of them”. Although this study involves different countries (Brazil, The Netherlands, Argentina and United States, the analysis was carried in the same market in all countries (soybeans), but at distinct levels (soybeans FOB Brazil, Argentina and United States and CIF Rotterdam), thus the above definition applies.

[15] CIF prices: all expenses, including freight and insurance, up to the port of destiny are at sellers’ account.

[16] FOB prices: all expenses up to the delivery of the merchandise on board of the ship indicated by the buyer, in the port of origin, are at sellers’ expenses.

17 According to the United States Standards for Soybeans, the are two classes of soybeans, mixed and yellow soybeans, the latter being "that have yellow or green seed coats and which in cross section, are yellow or have a yellow tinge, and may include not more than 10.0 percent of soybeans of other colors". Soybeans may classified in grade requirements ranging from 1 to 4, according to minimum test weight per bushel, maximum percent limits of damaged kernels and maximum count limits of other materials. Number 1 meets the highest standards.

[18] According to BARROS (1990, p. 13) the concept of impulse elasticity measures the variation of “variable i over variable j as a rate of the impact experienced by variable j and the impact experienced by variable i. The word elasticity applies strictly when variables are measured in logarithmic scale”.

[19] Refers to the notation of the variables used along the paper.

4.2 – Methods

The order of integration of the variables was determined with the Augmented Dickey Fuller (ADF) test in accordance with DICEKY and FULLER (1981 and 1979). Critical values were obtained in MACKINNON (1991) while critical values for the joint tests are from DICKEY and FULLER (1981).

It was used Granger's causality tests according GRANGER (1969) to verify the causality direction among the variables.

The cointegration test[20] aimed at detecting long-term relationship among the variables. This paper used the Johansen cointegration test following JOHANSEN and JUSELIUS (1990), with critical values from OSTERWALD-LENUM (1992). An Error Correction Model (ECM) was also estimated. According to BANERJEE (1993, p. 139), the ECM allows for a link between the short and long term dynamics as it provides a methodology for modeling both in levels and in differences. Consequently, the ECM models simultaneously the short-term dynamics of adjustment (variations) and the long term (levels). Other methodological aspects that need to be emphasized include the incorporation of restrictions over the short term parameters (α) and long term parameters (β) in the vector error correction model, besides the usage of a decomposition of the forecast errors and impulse-response function for the analysis of the price transmission dynamics over time in the international market for soybeans, and also exogeneity tests to check whether variables react to short term deviations from equilibrium.

5 – ANALYSIS OF THE RESULTS

Unit root testing requires choosing the number of lags in each test so as to eliminate autocorrelation in the residuals. Based on the Schwarz Information Criterion, variable *LROT* in level entered with two lags while in difference there was no need for lags. For *LBR* in level, the Criterion showed its minimum value for a second order autoregressive model, so that two lags were used in the respective unit root test. Variable *LBR* in differences showed a first-order autoregressive model. Thus, the unit root test procedure was carried out with only one lag. For variable *LARG* in level, the information criterion reached its lower value for an autoregressive model of order 2. In this case, we used two lags in the unit root test. For the same variable, differentiated of order one, the information criterion reached its lower value for a moving average model of order 1[21], leading to the use of the data dependent method in the identification of the number of lags[22], which resulted in six lags. Finally, for variable *LUSA* in level, it was identified an autoregressive model of order 2, as shown by the result of the information criterion, implying two lags. For variable *LUSA* differentiated, the information criterion detected a moving average model of order one, leading to the use of the data dependent method as it was the case with variable *LARG* differentiated. Thus, results from the *data dependent* method led to the inclusion of five lags in its respective unit root test (see Table 1).

Table 1. – Determination of the number of lags for the ADF-type unit root tests to variables *LBR*, *LROT*, *LARG* and *LUSA* (sample October 1995 – October 2003) using the Schwarz Information Criterion (BIC)

Variable	Minimum Schwarz Criterion Value	Lags effectively used in the ARMA model
LBR[1]	BIC(2,0) = -6.05439	Two lags
∇ *LPBR*[2]	BIC(1,0) = -6.07525	One lag
LROT [1]	BIC(2,0) = -6.0484	Two lags
∇ *LROT* [2]	BIC(0,0) = -6.0789	No lag
LARG [1]	BIC(2,0) = -5.86409	Two lags
∇ *LARG* [2]	BIC(0,1) = -5.88669	Six lags [3]
LUSA [1]	BIC(2,0) = -5.95719	Two lags

[20] The definition of cointegration can be found in ENGLE and GRANGER (1991).

[21] A moving average model of small-order can be written in the form of an autoregressive model of high order. Given that data are of monthly frequence, we used initially twelve lags.

[22] Details on the data dependent method are presented in PERRON (1994).

∇ *LUSA* [2]	BIC(0,1) = -5.96695	Five lags [3]

[1] Variable in level, [2] Variable in difference.[3] D*ata dependent method* as described by PERRON (1994).
Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

The variables presented unit root when tested in levels but became stationary in differences considering all the three significance levels that were adopted for each statistics. Table 2 shows that the first order differentiation produces stationary patterns for all variables, thus they are all difference-stationary, integrated of order one [$I(1)$].

Table 2. – Results from Augmented Dickey Fuller (ADF)[1] unit root tests to variables *LBR, LROT*, *LARG* and *LUSA*, October 1995 to October 2003.

Variable	τ_τ	ϕ_3	τ_μ	ϕ_1	τ	Order of Integration
LBR [2]	0.17	1.72	-1.10	0.66	0.28	I(1)
∇ *LBR*[3]	-5.61 [4]	16.12 [4]	-5.29 [4]	14.05 [4]	-5.32 [4]	I(0)
LROT [2]	-0.25	1.44	-1.33	0.88	0.00	I(1)
∇ *LROT* [3]	-8.12 [4]	33.03 [4]	-7.96 [4]	31.66 [4]	-8.00 [4]	I(0)
LARG [2]	-0.20	1.33	-1.22	0.75	0.03	I(1)
∇ *LARG* [3]	-3.69 [5]	7.18 [5]	-3.12 [5]	4.91 [5]	-3.15 [5]	I(0)
LUSA [1]	-0.36	1.18	-1.27	0.81	-0.03	I(1)
∇ *LUSA* [2]	-4.70 [4]	11.52 [4]	-4.16 [4]	8.66 [4]	-4.18 [4]	I(0)

[1] Critical values for τ_τ, τ_μ e τ obtained as described by MACKINNON (1991), respectively –3.1539, -2,5826 e –1.6175 at 10.0% level, and –3.4566, -2.8915 and –1.94345 at 5.0% level, and –4.056, -3.4993 and –2.5873 at 1.0%, level, while critical values for ϕ_3 e ϕ_1 were directly obtained from DICKEY and FULLER (1981) and are equal to 5.47 and 3.86 at 10.0% level, and 6.49 and 4.71 at 5.0% level, and 8.73 and 6.70 at 1.0% level. [2] Variable in level, [3] Variable in difference. [4] Significant at 1.0% level,[5] Significant at 5.0% level, [6] Significant at 10.0% level.
Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

Prior to proceeding with causality tests, it was necessary to choose the number of lags. In this exercise it was used the Akaike information criterion as presented by AKAIKE (1976). The minimum value was obtained with an autoregressive model of order 2. There were used two lags in the causality test.

Causality tests showed that the null hypothesis (H0), *LBR does not cause LROT, LARG e LUSA*, cannot be rejected, meaning that the probability of incurring in Error Type I (to reject the null hypothesis when it is true) is below the level of significance of 5%, thus the soybean price in Brazil does not influence the behavior of the international soybean price, which is represented by Rotterdam, Argentina and United States prices. This result came in line with expectations, as several studies show that Brazil, although important producer and exporter of soybeans, behaves as a price taker in the world market. In its turn, the null hypothesis that the international price for soybeans, here represented by variables *LROT*, *LARG* e *LUSA does not cause LBR* is rejected, as the probability of incurring in Error Type I (to reject H0 when it is true) lies much higher than the significance level of 10,0%, more precisely 70,57%. Again as expected the international price of the soybean influences the price of this commodity in Brazil, meaning the direction of causality is unidirectional. This result is in line with several studies that evaluated this price relationship (Table 3).

Table 3 – Granger Causality Test results for variables *LBR, LROT*, LARG and LUSA, October 1995 to October 2003.

Test	Null hypothesis	χ^2 test	Degrees of freedom	Probability
1	*LBR does not cause LROT, LARG, LUSA*	13.79	6	0.0321
2	*LROT*, *LARG*, *LUSA does not cause LBR*	3.79	6	0.7057

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

Another important procedure before the cointegration test is the determination of which of the five cases presented by JOHANSEN and JOSELIUS (1990) and JOSELIUS (1995) should be adopted. The underlying reason is that when estimating the Vector Error Correction Model (VEC) from the Vector Autoregressive Model (VAR) the deterministic terms of the VEC may differ from those of the VAR. More precisely, when a deterministic cointegration relationship exists, the deterministic terms of the VAR model will not be present in the VEC. On the other hand, should the relationship be a stochastic cointegration, the deterministic terms appear in the VEC embedded in the error correction term or as an independent term in the VEC. Case 2 was chosen through visual inspection.

The Johansen cointegration test found that at the 5.0% level, one cannot reject the null hypothesis that there is no cointegration vector against the alternative hypothesis that there is one cointegration vector. The following step was to test the null hypothesis that there is one cointegration vector against the alternative hypothesis that there are at least two cointegration vectors. Again, the null hypothesis was rejected at the 5.0% level, once the calculated value of the λ_{trace} statistics is higher than the respective critical value (see Table 4). Thus, we conclude that there are two cointegration vectors. Other relevant aspect is that the eigenvalues lie inside the unit circle. According to JOHANSEN (1995), this means that the type of non-stationarity presented by these variables in level can easily be removed through a difference operator, i.e., confirming that the variables are difference-stationary (DS). Since the number of cointegration vectors is smaller than the number of variables (reduced rank), the chosen model is the Error Correction Model (VEC) instead of the Vector Autoregressive (VAR)[23].

Table 4 – Results of Case 2 for the Johansen cointegration test on statistics λ_{trace}, variables *LBR, LROT*, *LARG* and *LUSA*, October 1995 to October 2003.

H_0:	H_1	Eigenvalue	λ_{trace}	Critical Value [1]	Error Correction Term	Error Correction Model
Rank = r	Rank > r				Constant	Constant
0	0	0.3517	85.67 *	53.42		
1	1	0.2601	44.50 *	34.80		
2	2	0.1394	15.88	19.99		
3	3	0.0168	1.61	9.13		

[1] Critical value provided by SAS at 5.0% level.
Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

When the analysis of the system involves more than two time series there may exist more than two cointegration vectors. According to DICKEY, JANSEN, THORNTON (1994, p.22,) cointegrating "vectors can be thought of as representing constraints that an economic system imposes on the movement of the variables in the system in the long-run. Consequently, the more cointegrating vectors there are, the "more stable" the system. Other things being the same, it is desirable for an economic system to be stationary in many directions as possible".

Results obtained from the estimation of the Vector Error Correction Model both in the short- and the long run were incompatible to economic theory *a priori* expectations, so that such results will not be presented here.

With the purpose of confirmation of the Law of One Price in the international market for soybeans it was imposed the restriction that long term parameters (β)[24] for variables *LROT*, *LARG* and *LUSA* are equal to

[23] More details on inter-relation between results of cointegration tests and the use of VAR models (in level or differences) or VEC can be found in HARRIS (1995).
[24] A discussion on restrictions to short and long term parameters is presented in JOHANSEN (1995) and HARRIS (1995).

one. The vector error correction model was estimated again, this time with the restrictions that *LROT*, *LARG* and *LUSA* are equal to one. Because of such imposition, the H matrix becomes the following (Box 1):

Box 1 – *H* matrix with restrictions imposed on β parameters

Variable	Vector 1	Vector 2
LPBR	1	1
LROT	-1	1
LARG	-1	1
LUSA	-1	1

Source: Adapted from Statistical Analysis Software (SAS).

Once the restrictions that $\beta_{11} = -\beta_{21} = -\beta_{31} = -\beta_{41}$ were imposed, the vector error correction model was re-estimated, taking into account the restrictions on the long-term parameters that affect the short-term elements α.

From the results of the cointegration test it is also possible to check whether the signs of the coefficients are in line with the prediction of economic theory. This is done with the analysis of the coefficients of the variables of the first cointegration equation normalized. Normalization was based on the estimate of the *LBR* coefficient, thus the coefficient estimate was set at one. The coefficient for *LBR* is the system's endogenous variable while *LROT*, *LARG* and *LUSA* are exogenous. In such a context, the analysis of the coefficient estimates for *LROT*, *LARG* and *LUSA* in the right column of Table 5 must be carried with inverted signs, since the normalized cointegration equation has all variables on the same side.

Table 5 shows that when the restrictions on the long-term β parameters are imposed, short term deviation from equilibrium represented by parameters α are eliminated at a speed of 26.16% in each period for Brazil against 31.27% of the soybean prices in Argentina. Apparently such results were able to detect the characteristics of each market. In spite of Argentina, Brazil shows a significant domestic market for soybeans. Brazil is the largest world exporter of chicken and soybeans are one of the main components of their feeding. As the domestic market is of significant size, one should expect that any short term deviation from equilibrium that may come from the international market would take more time to fade away in the Brazilian market as compared to the Argentine market. The latter being more linked to developments in the international market, it implies a higher sensibility of domestic prices against transitory deviation from equilibrium in external prices due to the absence of a significant domestic market. As compared to Brazil and Argentina, the Rotterdam response to deviations is even slower, at around 21.0% per period. Soybean prices in Rotterdam tend to react more slowly against prices of this commodity in Brazil and Argentina, for, differently of Brazil and Argentina which are price takers, Rotterdam is a price maker. Soy meal is mainly used as one of the key inputs for animal feeding, and the European Union is one of the world's largest producers of pork and cow meat. It is necessary to observe that, inversely to what occurs in Brazil, European cattle is fed with ration[25] and confined because of scarcity of land and feeding. In Brazil, cattle are grown in large properties fed with pasture. MACHADO; MARGARIDO (2003) used the X12 method and found that prices in Rotterdam show the smallest seasonal variation relatively to the other three markets (Brazil, Argentina and the United States), as a result of the alternation of harvesting seasons between the two hemispheres. The European Union faces a continuous supply the whole year, alternating from the northern hemisphere or the southern hemisphere harvesting seasons and that induce a pattern of stability in its domestic prices in response to short term deviations in external prices. Finally, the country with the smallest degree of sensibility of correction of short-term deviations was the United States, with a speed of only 2.5% each period. This relatively small speed of adjustment is related to the fact that, besides being the largest producer and exporter, the United States has a significant domestic market, being a price maker together with Rotterdam. These reasons lie

[25] The main components of animal ration are corn as source of carbohydrate and soy meal as a relevant source of proteins.

behind the slow fade away of short term deviations in the United States, meaning that its prices are less sensitive to short term variations of international prices relative to other markets.

Table 5– Estimates of short and long term coefficients of the Vector Error Correction model (VEC) with restrictions on parameters β, variables LBR, *LROT*, *LARG* and *LUSA*, Oct 1995 to Oct 2003.

Variables	Estimates of the short term adjustment coefficients (α) taking into account restrictions on β	Restrictions on long term parameters (β)
LBR	-0.26158	1
LROT	0.21003	-0.33300
LARG	-0.31273	-0.33300
LUSA	0.02500	-0.33300

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

The estimates of the long term parameter (β), with restrictions imposed on them, show that in the long run the Law of the One Price works in the international market for soybeans, for it was shown that the sum of these three estimated coefficients ($\beta_{21} + \beta_{31} + \beta_{41}$) equals unity (0.3333 + 0.3333 + 0.3333)[26]. Another relevant aspect of the results is the confirmation by this study that all long term estimated coefficients have the same value, thus confirming the working of the arbitrage process in the international market for soybeans. Thus, variations in the international price of soybeans tend to be fully transmitted to prices in Brazil in the long run, so that the elasticity of price transmission is unity. Finally, to confirm that the imposed restrictions are significant, i.e., to validate the Law of One Price, it was carried a joint χ^2 test to the set of restrictions, β taking unity value for *LROT*, *LARG* and *LUSA*.

Results of the χ^2 test show that the probability of Type I Error, of not rejecting the null hypothesis the coefficient of the restricted parameters are jointly significant ($\beta_{11} = \beta_{21} = \beta_{31} = \beta_{41} = 1$), when it is actually false, lies below 1%. The null hypothesis is accepted against the alternative of hypothesis that the restricted parameters are not significant ($\beta_{11} \neq 1$, $\beta_{21} \neq 1$, $\beta_{31} \neq 1$, $\beta_{41} \neq 1$), as presented in Table 6. In economic terms, this means that the Law of One Price holds in the international market for soybeans, since variations in the international price are fully transmitted to domestic prices in Brazil, having unity elasticity of transmission in line with the prediction of economic theory.

Table 6 –Johansen cointegration test with restriction for one cointegration vector.

Rank	Eigenvalue on Restrict	Eigenvalue	χ^2	Degrees of freedom	Probability value
1	0.1332	0.3376	41.46	4	<.0001

* Significant at 1.0% level.
Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

We carried tests for exogeneity on the short term parameters α, for confirming whether soybean prices in Brazil are influenced by external prices, thus identifying whether variables react to changes in the long run equilibrium. The analysis is conducted between the "dependent" variable and each one of the remaining "independent" variables. On the pairs *LBR* and *LROT* the null hypothesis that the soybean

[26] Here is important to notice that for VEC and VAR models there is not the traditional distinction between dependent and independent variables. These types of models have all variables as endogenous, thus the variable to be considered dependent needs to be normalized, as its coefficient is fixed as being equal to unity. Because of such normalization, the remaining variables are shifted to the right side of the equation, taking the role of independent variables. Thus, the signal of the restricted coefficient are inverted, so as that the interpretation in Table 5 requires inverting the signals. A plus sign must be read as a minus and conversely.

prices in Rotterdam are weakly exogenous, or that soybean prices in Brazil do not react to variations in originated in Rotterdam, was rejected, with 42.76% probability of incurring in Error Type I (not to reject the null hypothesis when it is false). Therefore, the null hypothesis was rejected in favor of the alternative hypothesis, thus the variable soybean price in Rotterdam is not weakly exogenous and prices in Brazil respond (are influenced) by price variations originated in Rotterdam. For Argentina we found the opposite. More precisely, the null hypothesis that soybean prices in Brazil do not react to price variations in Argentina, or that these are weakly exogenous cannot be rejected, for the probability of incurring in Error Type I is below the significance level of 1.0%, at around 0.031%. Thus price variations originated in Argentina does not affect prices in Brazil, the first being weakly exogenous relatively to prices in the Brazilian market. Finally, a result similar to that for Rotterdam was found on soybean prices in the United States, or that these are not weakly exogenous, and consequently, prices in Brazil react to price variations in the US market (see Table 7). The results that were found in exogeneity tests detected structural features of the international market for soybeans. Both Rotterdam and Chicago are price makers in the world market for soybeans, but it is worthwhile noticing that variation in prices in Rotterdam are more significant than variations in US prices on prices in Brazil. This happens because the European Union is a net importer of soybeans, basically used as animal feed, and also because it is the main destiny of soybeans exported by Brazil. One should then expect domestic soybean prices in Brazil to be influenced by variations in those two markets. On Argentina, the result also reflects the characteristics of this market, as Argentina, like Brazil, is a price taker in the world market.

Table 7 – Exogeneity test on short term parameters (α) for variables *LBR*, *LROT*, *LARG and LUSA*, October 1995 to October 2003

Variable	χ^2	Degrees of freedom	Probability value
LBR	0.64	2	0.7263
LROT	1.70	2	0.4276
LARG	11.57	2	0.0031
LUSA	2.72	2	0.2573

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

Table 8 presents the results on the variance decomposition of forecast errors on the four relevant variables. According to MARGARIDO (2000, p. 132-133), “the variance decomposition of forecast errors gives the dynamic behavior shown by economic variables. In particular, this procedure separates the variance of the forecast errors for each variable in components that may be attributable to each of the remaining endogenous variables, or it states in percentage the effect of an unanticipated shock on the remaining variables of the system”. The second column of Table 8 shows the periods expressed in months. In this paper, it is assumed that an unanticipated shock on any of the variables lasts for a maximum period of twelve months.

For variable LBR, the third column shows the percentage of variance of forecast errors against unanticipated shocks on this variable, or it evaluates the effect in time of an unanticipated shock on LBR over itself. The 4th, 5th and 6th columns detail the percentage of variance of forecast errors of LBR that can be attributed to variations in *LROT*, *LARG* e *LUSA*, respectively. From Table 8 we can see that 12 months after an unanticipated shock over LBR, there remains only 1.828% of the variance of forecast errors of LBR are caused by variables *LROT* (1.123%), *LARG* (0.221%) and *LUSA* (0.484%), while the remaining 98.17% are caused by itself. Thus, unanticipated shocks on Brazil grains prices are heavily influenced by itself, while the share of influence of the remaining international soybean prices are almost negligible, confirming the hypothesis that Brazil is a price taker in the international market for soybeans.

Table 8 – Results of the decomposition of variance of the forecast errors in percentage for variables *LBR*, *LROT*, *LARG* and *LUSA*, October 1995 to October 2003

Variable	Lead	lbr	Lrot	larg	lusa
Lbr	1	1.00000	0	0	0

3	0.99553	0.00257	0.000408	0.00149
6	0.98781	0.00756	0.00184	0.00278
9	0.98327	0.01025	0.00223	0.00425
12	0.98173	0.01123	0.00221	0.00484

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

The results of the variance decomposition of forecast errors of *LROT* are that one month after an unanticipated shock on that variable, 43.677% is caused by itself and 56.323% comes from *LBR.* Twelve months after the initial shock the already significant influence of *LBR* rises further to 75.816%, while the participation of *LROT* falls down to 21.483%. The influences of *LARG* and *LUSA* are 0.03260% and 2.66800% respectively, summing up to a total of only 2.70% (Table 9).

As for Brazil, it is relevant to notice that prior to the Complementary Act #87, of September 13, 1996, known as Lei Kandir, there was a relative equilibrium between Brazilian exports of grains and soymeal. As noticed by MARGARIDO e TUROLLA (2003), after that Act the composition of Brazilian exports of the soy chain has changed towards a smaller quantity of soymeal and stronger sales of grain, as the latter and primary products in general were exempted from taxes. More specifically, "the exemption of primary and semi-elaborated products caused a change in incentives to export, as compared to the other elements of the value chain. This change may have caused faster expansion of the segments that benefited from the exemption, at least early after the Act came into force" (MARGARIDO; TUROLLA, 2003, p.11)[27]. The results then show that there is a close relationship between soybean prices in Rotterdam and Brazil, as opposite to the relationship between prices in Argentina and the United States. Twelve months after an unanticipated shock on the Rotterdam soybean price, 75.82% of the Rotterdam price change is explained only by the prices in Brazil. In addition, on Argentina and the United States, again the results detected the international trade structure of the soybeans market.

Twelve months after an unanticipated shock on Rotterdam prices, the percentage share of the decomposition of forecast errors on Argentina is just 0.0326%, quite low as compared to the Brazilian figure. This possibly reflects not only the fact that Argentine soybeans are genetically modified[28], thus subject to import restrictions by the European Union, but also that most of Argentine exports in the soy chain are soymeal, unlike Brazil. And the European Union has barriers on imports on soymeal, while grain is imported without tariffs. The main objective of this policy is the creation of income and employment via incentives for the aggregation of value to take place inside the Union. In its turn, the Argentine policy creates incentive for soymeal exports relatively to soybeans. According to FREITAS et al. (2001, p. 8-9), "Argentina sends 96.0% of its production to the external market and keeps only 3.5% for domestic consumption. That country has in place a mechanism for protection of the local industry called *Retenciones*, which is designed to keep the raw material in the country by taxing grain exports at a 3.5% rate. It assures that local processing takes place, consequently raising employment in this chain. The mechanism creates an incentive to oil and soymeal production, without creating similar incentives to consumption. On the other hand, soymeal exports benefit from the *Reintegro*, a rebate in domestic taxes in the range between 1.4% and 6.8%. The mechanism drives Argentine prices downwards, raising their share in the international market".

[27] The change in the composition of exports may be related to a process of "reprimarization" of the Brazilian economy as named by GONÇALVES (2001, p.13-14). According to the author, "the reprimarization may be viewed under two aspects. The first comes because of the loss of international competitiveness of manufactures and the gain in agricultural products exported by Brazil. The second expresses the change in the structure of exports into a higher relative share of agricultural products and a smaller share of manufactures".

[28] It is important to keep in mind that the European Union imposes restrictions on imports of genetically modified soybeans. Among the major exporters, only Brazil produces conventional soybeans, while US and Argentina use genetically modified seeds.

The decomposition of variance of forecast errors for an unanticipated shock on the soybeans in Rotterdam shows that after 12 months only 2.668% of the Rotterdam price is caused by US prices (Table 9). Such a weak influence apparently has two causes. First, as mentioned earlier, US soybeans are genetically modified, and consequently its imports are limited by the European Union. In addition, the major market for the US soybeans is composed by Southeast Asian countries, not the European Union, the latter being mostly supplied by Brazil. Possibly these two factors help explaining the weak relationship between European Union and United States prices.

Table 9 – Results of the variance decomposition of forecast errors in percentage for variables *LBR*, *LROT*, *LARG* and *LUSA*, Oct 1995 to Oct 2003

Variable	Lead	lbr	lrot	larg	lusa
Lrot	1	0.56323	0.43677	0	0
	3	0.68974	0.29941	0.000582	0.01026
	6	0.73836	0.23587	0.000282	0.0255
	9	0.75198	0.22129	0.000333	0.0264
	12	0.75816	0.21483	0.000326	0.02668

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

The decomposition of variance of forecast errors show that one month after the unanticipated shock in Argentine prices only 20.60% is due to itself while Brazil prices enter with 78.84%. Rotterdam and US prices show an almost negligible influence of 0.563% and 0.0% (Table 10). Twelve months after the shock, Brazil contributes with 95.68%, Rotterdam with 1.902% and 0.606% (Table 10).

Table 10 – Results of the variance decomposition of forecast errors in percentage for variables *LBR*, *LROT*, *LARG* and *LUSA*, October 1995 to October 2003

Variable	Lead	lbr	lrot	larg	lusa
larg	1	0.7884	0.00563	0.20597	0
	3	0.90427	0.00338	0.08823	0.00412
	6	0.93425	0.0173	0.03873	0.00973
	9	0.95023	0.01848	0.02433	0.00697
	12	0.95681	0.01902	0.01811	0.00606

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

Like in previous cases, the results shown here detected conditions that prevail in the international market for soybeans. The relevant influence of Brazil prices in the short term on Argentine prices is possibly related to geographical conditions. These countries being neighbors and more important, located in the same hemisphere, they have similar harvesting seasons. However, as pointed out by MACHADO; MARGARIDO (2004), there is a one-month lag between the two countries' crops, with the Brazilian harvest coming one month before that of Argentina. As the second largest world producer, this one-month lag may be part of the explanation of the results of this paper, since crop expectations in Brazil influence not only Brazil prices but may also have considerable impact on the prices in Argentina. Box 2 depicts the entry times of the harvests of the major producers and exporters, making clear that the US harvesting season is inverse against the southern hemisphere countries, which show a high degree of time coincidence.

Box 2.- Soybean Crop and Trade Time

	Jan.	Feb.	Mar.	Apr.	May	Jun.	Jul.	Aug.	Sep.	Oct.	Nov.	Dec.
USA	**H**	**H**	**H**						**H**	**H**	**H**	**H**
BR			**H**	**H**	**H**	**H**	**H**	**H**	**H**			
ARG				**H**	**H**	**H**	**H**	**H**	**H**	**H**		

H = harvest period
Source: MACHADO; MARGARIDO (2004).

As it was mentioned earlier, the low influence of Rotterdam on Argentina may be related to the fact that most of the Argentine exports are soymeal, not soybeans; and also that European Union tariffs and non-tariff barriers on genetically modified soybeans, just the variety produced by Argentina.

The key reasons for the weak influence of US prices on Argentina may be related to different entry periods of entry in the market, despite being price taker, as opposite to what occurs between Brazil and Argentina. Box 2 shows that the US harvest goes from September to March, while the Argentine harvest is from April to October. This may be behind the weak interaction between both prices in the short term.

One last task is to analyze effects of shocks of US prices on the remaining variables. One month after such a shock, the decomposition of variance shows that 37.655% is due to itself, 59.18% to Brazil's prices, 1.888% to Argentina and 1.276% to Rotterdam. Twelve months after the shock, the own effect of the US price falls to only 3.641%, while the Brazilian influence rises to 92.01%, Rotterdam goes to 4.151% and Argentina only 0.198% (Table 11).

The strong relationship between prices in US and Brazil is possibly related to the share of both countries in global production and exports. However, data in Box 2 may offer some additional explanation: in March, when the US season is finishing and the Brazilian soybeans start entering the market. In its turn, when the Brazil season is finishing, there begins the US period. By the end of the US harvest, stocks decline and agents turn their expectations to the Brazil harvest, which is starting to trade. Thus, prices stay highly connected.

In the short term, the low share of Rotterdam prices relative to US prices possibly comes from the fact that the main market for US soybeans is in Asia, while Rotterdam is basically supplied by Brazil. Against that background, the Rotterdam market may be viewed as residual for US producers, thus explaining its little contribution for the formation of US prices.

The little contribution of Argentina for US prices possibly reflects the facts already mentioned, namely that US is a price maker while Argentina and Brazil are price takers, and also that Argentina incentives soymeal exports relatively to soybeans. Additional explanation for this phenomenon may be found in the natural switch of harvesting seasons between the major producers. As the end of the US season coincides with the beginning of the Brazilian season, which in turn comes one month before the Argentine, expectations on the Brazil prices are key to price formation in the US low season. That comes in line with the conclusion that Brazilian prices contribute heavily to the formation of US prices while the effects of Argentine prices on US are attenuated.

Table 11 – Results of the decomposition of variance of forecast errors in percentage for variables *LBR*, *LROT*, *LARG* and *LUSA*, October 1995 to October 2003

Variable	Lead	lbr	lrot	larg	Lusa
lusa	1	0.5918	0.01888	0.01276	0.37655
	3	0.74806	0.05934	0.00547	0.18714
	6	0.86554	0.04968	0.00238	0.08239
	9	0.90401	0.04356	0.00199	0.05045
	12	0.92010	0.04151	0.00198	0.03641

Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

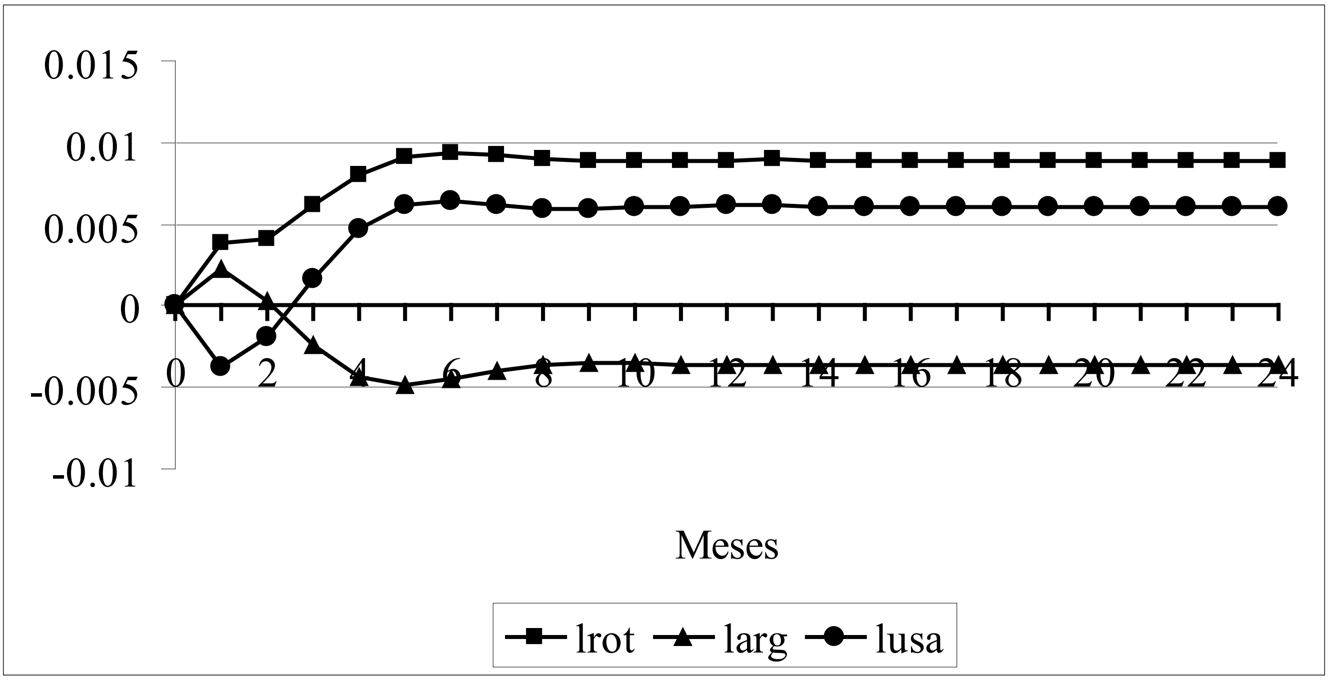
The following step was to calculate the impulse-response function. It shows how a shock in a variable influences the other variables of the system, thus helping to identify not only the size of the impact but also its time profile. As in the case of variance decomposition of forecast errors, it is possible to verify how impacts in each individual variable are transmitted to the remaining variables in the model. In this study, only the impacts of international prices on Brazilian prices are deemed to be relevant and are presented below.

Any unanticipated shock in Rotterdam induces an increase in Brazil prices. This upward path goes in four separate phases. Initially, Brazil prices increase rapidly up to one month after the shock. Between the 1st and the 2nd month prices keep increasing, however at a lower pace, reaching a maximum and stabilizing between the 5th/6th month after the initial shock (Figure 1). This time path apparently reflects that Rotterdam is a price maker and Brazil is a price taker in spite of being the second largest soybean producer and exporter. Another relevant aspect is that most of the Brazilian production goes to the European Union through its main gate, the Rotterdam port.

A shock in the US prices causes, in a first step, a fall in Brazilian prices up to the first month. From that point onwards, the path is inverted so that prices slowly move upwards up to the 5th/6th month, and they stabilize at that point similarly to the precedent case (Figure 1). Again, this is a consistent behavior as long as the US like Rotterdam are price makers. The initial price decline in Brazil as a response to a shock in US prices is possibly related to the different harvesting seasons that were earlier shown in Box 2. By the end of the US season in March, the Brazilian soybeans production starts entering in the market. Initially there is excess of supply in that moment, temporarily reducing prices in Brazil. Later on, the quantity supplied by US declines and only Brazil offers soybeans, since Argentina, although has similar harvesting period, offers more soymeal than grain. Thus Brazil's prices tend to increase up to the 5th month.

A shock in Argentine prices influence Brazil's in three steps. Brazilian prices increase up to the first month and then reverse downwards up to the point they stabilize by the 5th/6th month, inversely to the precedent cases. Since harvesting seasons are almost equivalent for both countries, a shock in the Argentine price coming from international demand, early in the southern hemisphere harvesting season, implies a situation in which the international market sees lack of supply. International demand shifts to the southern hemisphere. Initially, the initial scarcity causes Brazil prices to follow Argentine's. The second step is when quantities supplied by Brazil and Argentina tend to increase and Brazilian prices head downwards to reach a minimum at the 5th month after the shock and stabilize there.

Figure 1. Elasticity of the impulse-response function, effects of shocks of *LROT*, *LARG* and *LUSA* on *LBR*



Source: Basic data from OILSEEDS (oct. 1995- oct. 2003).

CONCLUSION

This paper investigates the price transmission in the world market for soybeans using time series econometrics models. The theoretical model developed by Mundlack and Larson (1992) is based on the Law of the One Price, which assumes price equalization across all local markets in the long run and allows deviations in the short run. The international market was characterized by three relevant soybean prices: Rotterdam Port, Argentina and the United States. The paper estimates the elasticity of transmission of these prices into soybean prices in Brazil. There were carried causality and cointegration tests in order to identify whether there is significant long-term relationship among these three variables. There was also calculated the impulse-response function and forecast error variance decomposition to analyze the transmission of variations over time in the international prices over Brazilian prices. An exogeneity test was also carried out so as to check whether the variables respond to short term deviations from equilibrium values. Results confirmed the Law of the One Price in the long run. In line with many studies, this paper showed that the speed of adjustment of prices in Brazil and Argentina can be seen as price takers as long as their response to shocks is slower than in the United States, the latter being a price maker.

A further research step may be to turn the international market more complete by incorporating the Chinese soybean prices. Although these were relevant to the study, a time series was not available.

An interesting conclusion was reached when the pattern of the impulse response functions was compared to the timing of crop and trade in Brazil, Argentina and the United States. The pattern of these functions is as follows:

- Brazilian prices respond to a shock in US prices with decay in the first month, and a rise above the initial level up to the $5^{th}/6^{th}$ month when it stabilizes.
- Brazilian prices respond to a shock in Rotterdam prices with a steady increase up to the $5^{th}/6^{th}$ month when it stabilizes, but the increase is more rapid in the first month
- Brazilian prices respond to a shock with an increase in the first month, and a decline below the initial level up to the $5^{th}/6^{th}$ month when it stabilizes. This pattern is opposite to the pattern of response to a shock in US prices.

These seasonal differences may help explaining the pattern of the response of Brazilian prices to shocks in the international market. Two interesting points can be made:

1. The response from shocks in the United States is opposite to the response from shocks in Argentina because harvest in the two hemispheres occurs in different periods.
2. The one-month lag between Brazilian and Argentine harvests may contribute to explain a turning point in the impulse-response function that occurs one month after the shock. Even the Rotterdam shock causes a more rapid increase of Brazilian prices in the first month than in the subsequent months.

This timing pattern was investigated in MACHADO; MARGARIDO (2004) and in this paper, and may deserve further research efforts.

REFERENCES

AKAIKE, H. Canonical correlations analysis of time series and the use of an information criterion. In: **Advances and case studies in system identification**. New York: Academic Press, 1976. p.27-96.

BANERJEE, Anindya *et al*. **Cointegration, error-correction, and the econometric analysis of non-stationary data**. New York: Oxford University Press, 1993. 329p. (Advanced Texts in Econometrics).

BARROS, Geraldo S.A.C. **Preços de commodities agrícolas e o câmbio flexível**. Piracicaba: ESALQ, 1990. 36p.

BARROS, Geraldo S.A.C; BURNQUIST, Heloisa L. Causalidade e transmissão de preços agrícolas entre níveis de atacado e varejo. In: Encontro Latino Americano da Econometric Society, 7, São Paulo, 1987. **Anais**... São Paulo, 1987, p 175-190.

CONAB. Exportações brasileiras principais produtos de origem agropecuária. Available at <www.conab.gov.br/download/indicadores/0205-balanca-exportacao.pdf> accessed on July 15, 2004.

DICKEY, David A.; FULLER, Wayne A. Likelihood ratio statistics for autoregressive time series with a unit root. **Econometrica**, v. 49, n.4, p.1057-1072, jul., 1981.

DICKEY, David A.; FULLER, Wayne A. Distribution of the estimators for autoregressive time series with unit root. **Journal of The American Statistical Association**, v. 74, n.366, p.427-431, Jun. 1979.

DICKEY, D.A.; JANSEN, D.W.; THORNTON, D.L. A primer on cointegration with application to money and income. In: RAO, B.B. **Cointegration for applied economists**. New York: ST. Martin's Press, 1994. p. 9-45.

ENGLE, Robert F.; GRANGER, C.W.J. **Long-run economic relationship**: readings in cointegration. New York: Oxford University Press, 1991. 301p. (Advanced texts in econometrics).

FREITAS, Silene M. ; MARGARIDO, Mario A. ; BARBOSA, Marisa Z. ; FRANCA, Terezinha J.F. Análise da dinâmica de transmissão de preços no mercado internacional de farelo de soja, 1990-99. **Agricultura em São Paulo**, 48(1); n.1, p. 1-20, 2001.

GONÇALVES, R. Competitividade internacional e integração regional: a hipótese da inserção regressiva. **Revista de Economia Contemporânea**. Rio de Janeiro, v.**5**, n. especial, p.13-34. 2001.

GRANGER, C.W.J. Investigating Causal Relations by Econometric Models and Cross-Spectral Methods. **Econometrica**, **37**, 1969, p.424-438.

HARRIS, Richard I.D. **Cointegration analysis in econometric modelling**. London: Prentice Hall, 1995. 176p.

JOHANSEN, Soren. **Likelihood-based inference in cointegrated vector auto-regressive models**. New York: Oxford University Press, 1995. 267p. (Advanced Texts in Econometrics).

JOHANSEN, Soren; JUSELIUS, Katarina. Maximum likelihood estimation and inference on cointegration with applications to the demand for money. **Oxford Bulletin of Economics and Statistics**, v. **52**, n. 2, p. 169-210, 1990.

KRUGMAN, Paul R.; OBSTFELD, Maurice. **International economics**: theory and policy. 4. ed., Massachusetts: Addison Wesley, 1997. 766p.

MACHADO, Eduardo L. ; MARGARIDO, Mario A. Seasonal price transmission in soybean international market: The case of Brazil and Argentina. Pesquisa e Debate, v.12, n.1 p.92-106, 2001.

MACHADO, Eduardo L.; MARGARIDO, Mario A. Evidences of seasonal price transmission in soybean international market. **Economia Aplicada**, v.8, n.1, p.127-141, jan.-mar. 2004.

MACKINNON, James G. Critical values for cointegration tests. In: ENGLE, R. F. ; GRANGER, W. J. **Long-run economic relationships**: readings in cointegration. New York: Oxford University Press, 1991. p.267-276.

MARGARIDO, Mario A.; SOUSA, Eduardo L.L.; BARBOSA, Marisa Z.; FREITAS, Silene M. Transmissão de preços no mercado internacional do grão de soja: uma aplicação da metodologia de séries temporais. In: CONGRESSO BRASILEIRO DE ECONOMIA E SOCIOLOGIA RURAL, 37., Foz do Iguaçu, 1999. Anais. Brasília: SOBER, 1999.

MARGARIDO, Mario A. ; SOUSA, Eduardo L.L. Formação de preços da soja no Brasil. **Agricultura em São Paulo**, v 45, n. 2, p. 52-61, 1998.

MARGARIDO, Mario A.; TUROLLA, Frederico A. Previsão de preços no mercado internacional de grão de soja. **Informações Econômicas.** São Paulo, v.**33**, n.1, jan. p.7-17. 2003.

MUNDLACK, Yair ; LARSON, Donald F. On the transmission of world agricultural prices. **The World Bank Economic Review**, v. 6, n. 1, p. 399-422, 1992.

OSTERWALD-LENUM, Michael. A note with quantiles of asymptotic distribution of the maximum likelihood cointegration rank test statistics, **Oxford Bulletin of Economics and Statistics**, v.54, n. 3, p.461-472. 1992.

PERRON, Pierre. Trend, unit root and structural change in macroeconomic time series. **In** RAO, B. Bhaskara. **Cointegration for the applied economist**. New York: ST. Martin’s Press, 1994. p.113-146.

PINO, Francisco A. e ROCHA, Marina B. Transmissão de preços de soja no Brasil. **Revista de Economia e Sociologia Rural**. Brasília, 32 (4):345-61, oct/dec. 1994.

USDA. Oilseeds: World Market and Trade. Washington: USDA, 1995/2003.

USDA. Oilseeds: World Market and Trade. Washington, Nov. 1997. p.13.

USDA. Oilseeds: World Market and Trade. Washington, Dec. 1999. p.11.

USDA. Oilseeds: World Market and Trade. Washington, Jan. 2003. p.9.

USDA. Oilseeds: World Market and Trade. Washington, Dec. 2003. p.9.