# Evidence of seasonal price transmission in the soybean international Market*

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## RESUMO

O objetivo do artigo é analisar o sistema de transmissão de preço no mercado internacional de soja. Inicialmente é determinado o sentido da causalidade do preço da soja no mercado internacional, utilizando as séries de preços do Chicago Board of Trade (CBOT), Roterdã, Brasil e Argentina, principais centros de produção e comercialização. Segundo, o comportamento sazonal das séries de preço foi avaliado usando-se o método X-12. Os resultados obtidos mostram que o sentido de causalidade entre as séries de preço analisadas é unidirecional. A análise dos índices sazonais indicou que o preço em Roterdã, no período de safra no hemisfério sul, é mais similar aos preços brasileiro e argentino. Por outro lado, os preços em Roterdã alinham-se às cotações da CBOT no período de colheita do hemisfério norte. Tal fato se deve ao consumo na União Européia, principal comprador mundial de soja, permanecer constante ao longo do ano. Por fim, conclui-se que o mercado internacional de soja apresenta um sistema de transmissão de preços baseado no comportamento sazonal.

**Palavras-chave**: soja, transmissão de preço sazonal e causalidade.

## ABSTRACT

The paper aims consists to analyze price transmission system in soybean international market. Initially is determined the price causality in soybean international market utilizing prices series from Chicago Board of Trade (CBOT), Rotterdam, Brazil and Argentina, production and commercialization main centers. Second, seasonal behavior of the price series was evaluated using X-12 method. The results obtained showed that causality directions among the analyzed price series are unidirectional. The seasonal behavior analysis indicated that Rotterdam price, at the crop time in South Hemisphere, is more similar to Brazilian and Argentina prices. On the other hand, Rotterdam prices align with CBOT quotation in North Hemisphere crop time. This result is due to the fact that soybean consumption in European Union, the largest international soybean buyer, to remains constant during the year. Last, the paper concludes that international soybean market presents a price transmission system based on seasonal behavior.

**Key words**: soybean, seasonal price transmission and causality.

**JEL classification**: C32, F31.

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## 1 Introduction

The traditional and dominant commercialization system in soybean international market has been strongly dependent of Chicago Board of Trade (*CBOT*) and Rotterdam Port. Prices are set in these centers, which dominate the commercialization of the product and influence the strategies of all chain agents.

Brazil and Argentina are two important players in soybean international market in terms of both production and exportation. Almost 70,00% of total Brazilian and Argentina exportation of soybean has been destined to European Union (*EU*) in last decade. Besides, the crop in these countries occurs when the international market is without product supply, due to the end of United States crop (the main world producer). Thus, the *CIF*[1] prices in Rotterdam are an important reference to *FOB*[2] prices in these countries. According to ABIOVE,[3] Brazil is the world second largest exporter, with something as 20 percent of world production.

This paper aims to bring new evidences and to analyze the seasonal behavior among soybean price in *CBOT*, *CIF* prices in Rotterdam and *FOB* prices in Brazil and Argentina. The main conjecture is that Brazilian and Argentina prices are more associated to Rotterdam than Chicago; due to the European Union is the main destination of soybean exported by both countries. The expected result is that the soybean price seasonal index in Rotterdam will have more adherence with producer countries seasonal index, in function of production time specificity. Furthermore, it is expected that Rotterdam seasonal index graphic trajectory has lower amplitude than the producer countries, due to supply stability in Rotterdam. In addiction, a complementary result of the first conclusion consists in find a possible seasonal behavior based in a price transmission system. This expected result was founded in preliminary study by Machado and Margarido (2001).

Aguiar and Barros (1991) and Neves (1993) use Granger causality test to determine the Brazilian soybean causality. The main conclusion is that Brazil doesn't set prices in international market. Pino and Rocha (1994) conclude that Brazilian soybean price is affected by *CBOT* variations, using ARIMA and Box-Jenkins transfer function models from 1985 to 1990. According Margarido and Sousa (1998), *CBOT* variations between

1 In CIF prices the seller is responsible for all expenses until goods final destination.

2 In FOB prices the seller is responsible for all expenses until goods shipment local.

3 Brazilian Association of Vegetable Oil Industries.

1990 and 1998 were immediately transmitted to Brazilian prices, using the ARIMA models developed by Haugh and Box (1977).This approach incorporates a causality test in transfer function model.

## 2 Materials and methods

### 2.1 Materials

The data set contains 120 observations for each series analyzed from January/1991 to December/00. The data for soybean quotations were obtained in Chicago Stock Market, *CHI*[4] (January/91 to October/98) and in Vegetal Oil Industry Brazilian Association, ABIOVE (November/98 to December/00). The Brazilian and Argentina Free on Board (*FOB*) prices and the Rotterdam (*ROT*) Port Cost Insurance and Freight (*CIF*) prices were founded in Oilseeds publication (several numbers)..

### 2.2 Methods

#### 2.2.1 X-12 method

The seasonal index of each series were obtained from Statistical Analysis Software (SAS, version 8.2), using the methodological framework developed by the U.S. Bureau of the Census and SAS Institute (1999).

According of Yaffe and Mcgee (2000, p. 66), "*The U.S Census has developed its X-12 program, which contains some innovations over the earlier X-11 and the 1988 update, X-11 ARIMA, developed by E.B. Dagum et al. Dagum had introduced X11 ARIMA to use back- and forecasting to reduce bias at the ends of the series. The new X-12 program contains more systematic and focused diagnostics for assessing the quality of seasonal adjustments.*" The new approach uses a wide variety of filters from which to permit ex-

4 Variable notation uses in this text. *CHI* refers Chicago Stock Market prices; *ROT* refers Rotterdam Port CIF prices; *BR* refers Brazilian FOB prices; *ARG* refers Argentina FOB prices.

5 Details about seasonality theory and its applications can be found in Pino *et al.* (1994).

tract trend and seasonal patterns and a second set of asymmetric filters to be used for the ends of the series.[5]

### 2.2.2 Unit root test

The integration order was determined using both Augmented Dickey-Fuller (ADF)[6] and Phillips-Perron (PP) unit root tests.[7] The critical values for unit root tests can be found in Mackinnon (1991).[8] In economic terms, the existence of unit root test in time series consists in an important procedure. For example, Alencar (1998, p. 171) shows that if the unit root hypothesis was true for an economic series, the random shocks would have a permanent effect. In this case the fluctuations will be not transitories.[9]

### 2.2.3 Causality test

Gujarati (1995) defines the causality concept as "*if variable x causes variable y, then changes in x should precede changes in y.*" A causality test relatively simple was proposed by Granger (1969). This test assumes the information relevant to the prediction of the variables is contained solely in the time series data on these variables. The two variables test estimated ($y$ e $x$) is represented by the regression below:

$$
\begin{aligned}
y_t &= \alpha_0 + \sum_{i=1}^{k} \alpha_i\, y_{t-i} + \sum_{i=1}^{k} \beta_i\, x_{t-i} + \varepsilon_{1t} \\
x_t &= \beta_0 + \sum_{i=1}^{k} \beta_i\, x_{t-i} + \sum_{i=1}^{k} \alpha_i\, y_{t-1} + \varepsilon_{2t}
\end{aligned}
\qquad (1)
$$

where the disturbances terms are uncorrelated. One important observation is that the number of lagged terms included in regression (1) can affect the direction of causality,

6 See Dickey and Fuller (1979, 1981).

7 See Phillips and Perron (1988). See Holden and Perman (1994) for a complementary reference.

8 Margarido and Anefalos (1999) provide a complete guide line about unit root test application.

9 See Nelson and Plosser (1982) for unit root importance in economic cycles.

because the Granger test is very sensitive to the number of lags used in the analysis. Gujarati (1995) distinguishes four possible results in regression (1):

1) Unidirectional causality from $x$ to $y$ exists if $\sum \beta_i \neq 0$ e $\sum \alpha_i = 0$;

2) Unidirectional causality from $y$ to $x$ is indicated if $\sum \alpha_i \neq 0$ e $\sum \beta_i = 0$;

3) Bilateral causality is suggested if $\sum \beta_i \neq 0$ e $\sum \alpha_i \neq 0$;

4) Independence or absent causality occurs if $\sum \beta_i = 0$ and $\sum \alpha_i = 0$.

The $t$ test is used to verify the individual statistical relevance of both $\beta'_i$ and $\alpha'_i$ parameters. The joint significance of the complete set of variables is tested using $F$ test.

## 3 Results

### 3.1 Unit root tests

The first step consisted in determination of the lags number for each series, using the Schwarz Information Criterion (BIC) as a decision parameter, except for the differenced variable *CHI*. In this case, it was necessary to utilize the data dependent method.[10] Table 1 summarizes the results obtained.

The second step consisted in the determination of integration order for each series. For *ROT* series, at level, the null hypothesis for unit root couldn't be rejected for all statistics at significance levels tested (1%, 5% and 10%). These results implied in the realization of ADF and PP tests again, but now with differenced variable. The null hypothesis was rejected to all statistics and significance levels tested. Therefore, *ROT* series was considered integrated of order one.

10 See Perron (1994) for a complete method description.

**Table 1**
**Lags Number Determination**

| Series | BIC minimum value | Effective Lags Number |
|---|---|---|
| $CHI$ [1] | BIC(1,1) = 4,778300 | 1 |
| $\nabla CHI$ [2] | BIC(0,1) = 4,785865 | 8 [3] |
| $BR$ [1] | BIC(1,0) = 4,933081 | 1 |
| $\nabla BR$ [2] | BIC(0,0) = 4,923930 | 0 |
| ARG [1] | BIC(1,0) = 4,996052 | 1 |
| $\nabla ARG$ [2] | BIC(0,0) = 4,991172 | 0 |
| $ROT$ [1] | BIC(1,0) = 4,718388 | 1 |
| $\nabla ROT$ [2] | BIC(0,0) = 4,703662 | 0 |

[1] Level, [2] Differenced, [3] Data dependent method started with 12 lags.

**Source:** Primary Data from Chicago Board of Trade (CBOT), Brazilian Vegetal Industry Association (ABIOVE) and Oilseeds (1991-2001).

The series *ARG* and *BR* had similar behavior. For each individual series, at level, the null hypothesis for unit root couldn't be rejected for four statistics ($\tau_\tau$, $\phi_3$ $\tau_\mu$ and $\phi_l$) in ADF test and for two statistics ($\tau_\tau$ and $\tau_\mu$) in PP test. Only for the $\tau$ statistic the two series rejected the null hypothesis. In function of these results, the autocorrelation function was used to choose the integration order of these series, in accordance with procedure describe in Box, Jenkins and Reinsel (1994), Mills (1990) and Vandaele (1993).[11]

The autocorrelation function graphic shows a slow and continuous decrease in the course of time. This shape consists in an indication of unit root presence, since both series have a strong "memory". So, the last result implied in the realization of ADF and PP tests again, but now with differenced variable. The new results show that null hypothesis was

11 It is necessary to observe that unit root tests have low power in relation to small samples, allowing high probability of type I error. Additionally, its results are very sensitive in relation to lag number used, constant and/or tendency introduction. According to Hatanaka (1998), cases which characteristic roots are inserted in the interval between 0.9 and 1.0, Dickey-Fuller test cannot distinguish an stationary difference process (DS), that is, a process that contains unit root, from a stationary tendency (TS) one, for a small sample.

rejected to all statistics at 1% significance level for both ADF and PP tests.[12] Therefore, the two series were considered integrated of order one. Similar procedure was used to determine the integration order of *CHI* series.

**Table 2**
**Augmented Dickey-Fuller Test Results**

| Series | $\tau_\tau$ | $\phi_3$ | $\tau_\mu$ | $\phi_1$ | $\tau$ | Integration Order |
|---|---|---|---|---|---|---|
| *CHI*[1] | -2,17 | 2,53 | -2,06 | 2,13 | -2,07[4] | I(1) |
| $\nabla$*CHI*[2] | -3,44[5] | 5,90[5] | -3,30[4] | 5,46[4] | -3,32[3] | I(0) |
| BR[1] | -1,70 | 1,68 | -1,62 | 1,32 | -1,63[5] | I(1) |
| $\nabla$BR[2] | -11,03[3] | 60,80[3] | -11,03[3] | 60,84[3] | -11,08[3w] | I(0) |
| *ARG*[1] | -1,94 | 2,04 | -1,89 | 1,79 | -1,90[5] | I(1) |
| $\nabla$*ARG*[2] | -10,02[3] | 50,20[3] | -10,04[3] | 50,36[3] | -10,08[3] | I(0) |
| *ROT*[1] | -1,65 | 1,60 | -1,50 | 1,15 | -1,51 | I(1) |
| $\nabla$*ROT*[2] | -10,42[3] | 54,32[3] | -10,42[3] | 54,31[3] | -10,47[3] | I(0) |

[1] Level , [2] Differenced , [3] Significant at 1,0% level, [4] Significant at 5,0% level, [5] Significant at 10,0% level.

**Source:** Primary Data from Chicago Board of Trade (CBOT), Brazilian Vegetal Industry Association (ABIOVE) and Oilseeds (1991-2001).

**Table 3**
**Phillips-Perron Test Results**

| Series | $\tau_\tau$ | $\tau_\mu$ | $\tau$ | Integration Order |
|---|---|---|---|---|
| *CHI*[1] | -1,89 | -1,76 | -1,76[5] | I(1) |
| $\nabla$*CHI*[2] | -8,32[3] | -8,34[3] | -8,38[3] | I(0) |
| BR[1] | -1,72 | -1,65 | -1,66[5] | I(1) |
| $\nabla$BR[2] | -11,03[3] | -11,03[3] | -11,08[3] | I(0) |
| *ARG*[1] | -1,91 | -1,86 | -1,86[5] | I(1) |
| $\nabla$*ARG*[2] | -10,02[3] | -10,04[3] | -10,08[3] | I(0) |
| *ROT*[1] | -1,63 | -1,48 | -1,63 | I(1) |
| $\nabla$*ROT*[2] | -10,42[3] | -10,42[3] | -10,47[3] | I(0) |

[1] Level, [2] Differenced, [3] Significant at 1,0% level, [4] Significant at 5,0% level, [5] Significant at 10,0% level.

**Source:** Primary Data from Chicago Board of Trade (CBOT), Brazilian Vegetal Industry Association (ABIOVE) and Oilseeds (1991-2001).

12 The visualization of autocorrelation function to each differenced variable presented strong downfall tendency in the first lags, showing stationary behavior of both variables.

### 3.2 Causality test

Causality tests results with stationary variables are presented in Table 4. The analysis of the relationship between Argentina e Brazil demonstrates that the null hypothesis of *ARG* not causing *BR* couldn't be rejected at 5% significance level. Analyzing the inverse way, the null hypothesis of *BR* not causing *ARG* was rejected. Therefore, the causality direction in this case is unidirectional. The results are coherent with the soybean market characteristic in each country. Brazil is the second biggest soybean producer and exporter in the world. Meanwhile, Brazilian consumption of soybean basic products is extremely high. This is due to the importance of Brazilian poultry production, since soybean is the main component of poultry feed composition.[13] On the other hand, Argentina is the third biggest soybean producer, but almost all production is exported due low internal consumption. So, Brazilian prices are less sensitive[14] to international price variation when compared with Argentina. This effect was verified in several studies, like Freitas *et al.* (2001) and Margarido *et al.* (1999).

Causality tests results showed coherence in the relation between Argentina's prices and Chicago's soybean quotations, as well as in the relation between Chicago's quotations and Brazilian soybean prices. Both situations had rejected the null hypothesis that soybean quotations from Chicago do not cause the price of the product in Argentina and Brazil at level 5.0% and 1.0%, respectively. On the other hand, the null hypothesis that soybean prices in Argentina and Brazil do not cause Chicago's quotations was not rejected. These results demonstrate coherence, that is, show that price is formed in Chicago, while Argentina and Brazil are price takers in the commodity international market. Thus, the causality directions are unidirectional for both cases.

In relation to the variables soybean quotations at Chicago's stock exchange and *CIF* prices in Rotterdam, the causality tests showed that causality direction is also unidirectional. The null hypothesis that soybean prices in Rotterdam do not cause the Chicago's product quotations was not rejected. Meanwhile, in the inverse direction, the null hypothesis that Chicago's soybean quotations do not cause Rotterdam's *CIF* prices was rejected at 5.0% significance level. The result that Rotterdam's prices do not influence Chicago's soybean quotations possibly reflects the fact that expressive share of United

13 Brazil is the second biggest poultry producer in the world.

14 Nevertheless, it is important to notice that both countries, while large soybean producers and exporters, are price takers in international market. Variations are related to the intensity which soybean international prices are transmitted to its respective domestic prices, as function of each country's own structural characteristics.

States' soybean production is destined to Asian markets, thus, attenuating the effect of demand variations from European markets, that is, United States' soybean prices are not so sensitive to Rotterdam's price variations, contrarily to Argentine and Brazilian situations, both with *EU* as main soybean consumer market. Inversely, Chicago's quotations influence Rotterdam's prices. Probably, as United States are the main soybean and its derivatives producer and exporter, price variations in this market are transferred to EU prices, which is the major international soy products consumer market.

Finally, it was verified that the relationship between Rotterdam's *CIF* prices and Argentina's and Brazilian *FOB* prices also presented unidirectional course. Both null hypothesis that Rotterdam's prices do not cause Argentina's nor Brazilian prices were rejected at 10.0% and 5.0% significance levels, respectively. Causality tests showed that the null hypothesis of Argentina's soybeans *FOB* prices not causing *CIF* prices was not rejected. Similar result was obtained for the test with Brazilian *FOB* prices and Rotterdam's CIF prices. These results confirm that Argentina and Brazil are price takers in the soybean international market, in accordance with several studies results.

**Table 4**
**Granger Causality Test Results**

| Null Hypothesis | $\chi^2$ Test | Freedom Degree | Probability |
|---|---|---|---|
| ARG not cause BR | 5,76 | 3 | 0,1237 |
| BR not cause ARG | 9,36 | 3 | 0,0249 * |
| ARG not cause CHI | 3,85 | 6 | 0,6965 |
| CHI not cause ARG | 12,29 | 6 | 0,0557 * |
| BR not cause CHI | 2,32 | 6 | 0,8882 |
| CHI not cause BR | 17,05 | 6 | 0,0091 * |
| ROT not cause CHI | 2,65 | 2 | 0,2661 |
| CHI not cause ROT | 6,43 | 2 | 0,0402 * |
| ARG not cause ROT | 0,00 | 1 | 0,9928 |
| ROT not cause ARG | 3,46 | 1 | 0,0627 * |
| BR not cause ROT | 5,07 | 6 | 0,5343 |
| ROT not cause BR | 14,28 | 6 | 0,0267 * |

* Significant.

**Source:** Primary Data from Chicago Board of Trade (CBOT), Brazilian Vegetal Industry Association (ABIOVE) and Oilseeds (1991-2001).

### 3.3 Seasonal indexes

Soybean harvest and commercialization occur in different periods in each one of the producer regions considered in the study. Figure 1 shows that in North Hemisphere (United States) soybean harvest and commercialization period is from September to March. On the other hand, South Hemisphere (Brazil and Argentina) the periods begin March/April and end in September/October, that is, present inverse behavior in relation to North Hemisphere.

**Figure 1**
**Soya Bean Crop and Trade Time**

| | J | F | M | A | M | J | J | A | S | O | N | D |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| EUA | | | | | | | | | | | | |
| Brazil | | | | | | | | | | | | |
| Argentina | | | | | | | | | | | | |

Source : Brazilian Vegetal Oils Industry Association (ABIOVE).

Given the each soybean producer characteristics; it is expected that Brazilian and Argentina seasonal indexes present similar behavior. Meanwhile, the inverse behavior in relation seasonal index is expected between Brazil/Argentina and United States as consequence of the different soybean harvest and commercialization periods.

It is also expected that Rotterdam seasonal soybean prices index present lower seasonal amplitude in comparison to the producer countries,[15] because Rotterdam, as entrance port for European Union market, is supplied during the all year, depending on crop period, by both United States and Brazilian and Argentina soybean production. So, it is expected that Rotterdam's seasonal soybean prices index are less inclined to sudden variations in comparison to the indexes from producer countries.

15 It is expected that seasonal amplitudes present higher variance in countries that produce and export soybean, due to its own cultivation characteristics. Thus, in harvest periods, there is a quantitative expansion on product supply and, consequently, lower prices; while in post-harvest period, there is a quantitative reduction on product supply, resulting in higher prices.

Another relevant aspect in relation to international soybean market is that only three countries produce almost all soybean commercialized worldwide. Thus, it is expected that, during the harvest period in North Hemisphere, Rotterdam seasonal prices index present more adherence in relation to Chicago seasonal prices index. On the other hand, a reverse situation occurs in relation to South Hemisphere harvest period. In this case, Rotterdam seasonal prices index is more close to Argentina and Brazilian seasonal prices indexes.

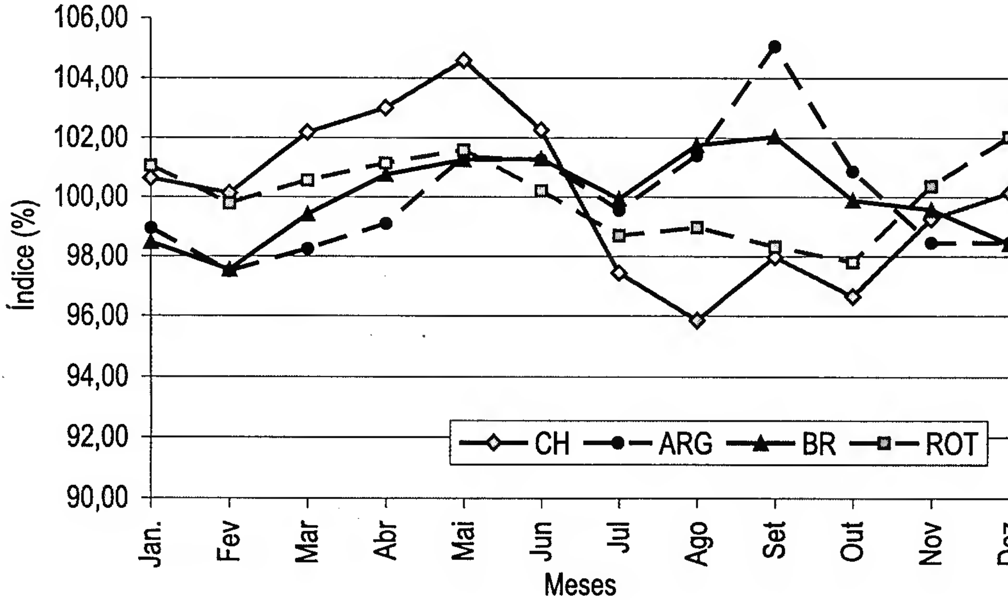
The seasonal index of *CBOT* quotation varied between 95,85 (August) and 104,58 (May), respectively the beginning of crop and the off-season period in the North Hemisphere, with amplitude coefficient 8,71%. These results are consistent because the prices during the crop are smaller than the off-season prices.

In Brazil, the seasonal index of *FOB* prices varied between 97,54 (February) and 102,05 (September), respectively the beginning of crop and the off-season period in the South Hemisphere, with amplitude coefficient 3,57%. Thus, as in the former case, the indexes seem to capture the market conditions.

Argentina's *FOB* prices reached minimum value of 97,48 in February and maximum value of 105,05 in September, with amplitude coefficient 7,47%.

The lower amplitude coefficient obtained for Brazil in relation to Argentina possibly reflects each countries soybean market conditions. As mentioned before, Brazil is a large soybean producer, exporter and consumer, due to the use of soybean as main input for poultry feed production, a situation that does not occur in Argentina. Thus, Argentina's prices tend to be more influenced by price variations in international soybean market than Brazilian domestic prices, what justifies the Brazilian wider amplitude coefficient in relation to Argentina.

**Figure 2**
**Seasonal Indexes**




Source: Primary Data from Chicago Board of Trade (CBOT), Brazilian Vegetal Oils Industry Association (ABIOVE) and Oilseeds (1991/2001).

The seasonal index of Rotterdam's *CIF* prices varied from a minimum of 97,80% in October to a maximum of 102,04% in December, with amplitude coefficient 4,24%.

The seasonal amplitude indexes results show that Rotterdam prices present low amplitude variation, presenting an amplitude coefficient higher only in relation to Brazilian coefficient. Probably due to the stability of soybean supply to the *EU* along the year, since during the South Hemisphere's off-season period the *EU* market is supplied by the North American crop; and, in the North Hemisphere's off-season period, the *EU* market is supplied by the South American crop.

In Figure 2, it is noticeable that during the harvest period in South Hemisphere, Rotterdam seasonal prices index are more close to Brazilian and Argentina seasonal prices indexes than in relation to Chicago's seasonal prices index, as it was expected. The situation is reverse during the harvest period in North Hemisphere, that is, Rotterdam seasonal prices index are more adherent to Chicago's seasonal prices index.

## 4 Conclusion

The traditional and dominant commercialization system in soybean international market has been strongly dependent of Chicago and Rotterdam Port. It's a common belief that prices from *CBOT* dominate the commercialization of the product and influence the strategies of all chain agents. Meanwhile, the results founded to confirm the dependency of both Brazilian and Argentina prices to foreign market and that Rotterdam is more important than Chicago to price formation in both countries.

There is a strong dependence of the Brazilian and Argentina *FOB* prices with the CIF prices in Rotterdam, differently of United States prices, that are set within *CBOT*. Margarido *et al.* (1999) and Machado and Margarido (2001) had obtained similar results to the Granger causality test that confirms the dependency of both Brazilian and Argentina prices to foreign market.

Other important result is that the amplitude of seasonal standard is more accentuated in United States off-season period in Brazil and Argentina. On the other hand, the seasonal standard in Rotterdam is less accentuated than another series, due to the fact that supply in European Union is constant during all year. These results are according to the expected for this market, due to United States, Brazil and Argentina crops occurs in distinct periods of the year.

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