Analysis of Soybean Price Transmission between Major Brazilian Producing Areas: A Co-integration and Causality Approach
Carlos Alberto Gonçalves da Silva¹, Salatiel Turra², Filipe Higino Dias de Souza³ & Léo da Rocha Ferreira⁴

Abstract
Soybean production plays a significant role in both the Brazilian and in the world economy. A relevant issue to be observed in agribusiness management is related to its price transmission. The main objective of this paper aims to analyze the price transmission process among major Brazilian soybean producing areas: States of Mato Grosso, Paraná, Rio Grande do Sul and Goiás, covering the period from January 2008 to December 2015. In order to fulfill this objective a bivariate co-integration test for soybean prices, as well as the auto-regression vector model (VAR) for the Granger causality analysis and function Impulse response. The results showed that there is price transmission between all the states analyzed, except for the State of Goiás.

Key words: soybean, price transmission, auto-regression vector model, Brazilian agriculture

Available online
www.bmdynamics.com
ISSN: 2047-7031

INTRODUCTION
Soybean production plays an important role in the Brazilian and the world economy in view of its multiple uses: protein generation for animal consumption as well as oil production for human consumption. Additionally, it is one of the major commodities of Brazilians exports.

According to Dall’Agnol; Lazarotto; Hirakuri, (2010), the soybean industrial complex presents a relevant performance both in national supply and demand levels and, in a particular way, promoting economic growth in several regions of the country. This fact is due to its continuous growing production since the half of the XX Century.

The notorious evolution of world production becomes more obvious, according to data of the Instituto Mato Grossense de Economia Agropecuária (IMEA), 2015, when we compare the 19.1 million tons harvested in 1964/65 with the 95.7 million tons harvested in 2014/15.

This growth was a result of production area enlargement, as well as productivity growth by harvested hectare, motivated by technological innovations in the genetics of the commodity and by the generated knowledge transferred to farmers. In this way, this national mark began in the 1960 decade, gaining significant proportion in the seventies (Barreto, 2011).

Some regions of the country became more expressive in soybean cultivation in view of their present soil and climatic conditions, because of their more efficient logistic infrastructure (production distribution) and by the presence of research nets directed to specific crops. The South Region and the State of São Paulo were the major regions able to exceed in this matter.

Up to the 1980 decade more than 80% of the soybean national production was concentrated in the three states of the South. However, other regions with similar climatic conditions to the South started to detach themselves by soybean cultivation. According to data of the Companhia Nacional de Abastecimento – CONAB (2015); this growth can be attributed to good climatic conditions plus its high technological level.

On recent crop periods (2011/12; 2012/13; 2013/14 and 2014/15), states with more production concentration of the commodity were: Mato Grosso, Paraná, Rio Grande do Sul and Goiás, respectively.

Major progress in production is a result of technological development, which may result in economic efficiency only if these production regions act in an integrated form so that their eventual production variations can be reflected on local prices. In this way, by transmitting desired supply movement signs in the short run and production adjustments in the medium and long run (Bittencourt; Barros, 1996). However, the price received by producers is in fact the major economic incentive in order to increase

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production. In this context, a relevant question to be observed is the one related to the identification of price transmission elasticity among major soybean production states. Therefore, according to Hoffmann (1998 *apud* Melo *et al.*, 2008), it must be considered that agricultural product prices show cyclical variations along different seasons of the year. Besides, other factors such as: the nominal variability of important input prices (seeds, fertilizers and herbicides) and climatic and soil conditions also generate variability in the soybean price.

Nevertheless, even if elastic responses are admitted integrated markets must present price convergence dynamic characterizing long run equilibrium. Therefore, the Law of One Price confirmation consists in a fundamental theoretical element to postulate market integration, as well as, a necessary presuppose so that price transmission may occur.

Besides implying the importance of market integration it is interesting to notice that the majority of the research already done, in general, has tried to identify price transmission only in internal and external markets but not in those regions of only one national market. Therefore, this research has the proposition of analyzing domestic prices with the objective to verify if the Law of One Price can be applied among these regional markets, *i.e.*, if price variations in one market can be transmitted to other markets.

Besides this introduction, in order to answer these questions about determinants of the price transmission elasticity analysis, this article is divided into five sections. The next section, presents the main objective of this paper. In the following section, a literature review about: (3.1) World Panorama; (3.2) National Panorama; (3.3) Price Transmission; and (3.4) The Law of One Price; with emphasis on soybean is presented. In the fourth section the methodology and the data used in the research are detailed. The fifth section presents the results of the estimated model. Finally, the sixth and last section presents its final considerations.

Given the importance of price causality in the agribusiness management the main objective of this paper is to examine price transmission and its casual relationship among four major Brazilian soybean production regions, from the January of 2008 to the December of 2015 period. Specifically, the objective of this paper is to verify if:

- The Law of One Price, that is, if soybean price variation in one state can be transmitted to other states; and
- Evaluate if the analyzed states are integrated.

**LITERATURE REVIEW**

**World Panorama**

The importance of soybean as a major commodity, among other things, is due to its diversified usability. It has the capacity and conditions of producing protein and oil for animal and human consumption, respectively. Additionally, it is one of the major commodities of Brazilian exports. The evolution of world soybean production was considerable in the second half of the last century. According to United States Department of Agriculture (USDA, 2016), there was an increase in production from 19.1 million tons harvested in 1964/65 to 320.49 harvested in 2015/16. Hence, soybean became the major crop in world commerce, even substituting other oils of vegetable or animal origin (SEAB/DERAL, 2014). The major world justifications for this commodity expansion are directly related to its production and productivity by hectare, which was continuously improved by the related generation of scientific knowledge and passed on to farmers.

According to the USDA, the United States was the largest world soybean producer in 2015/16, (SEAB/DERAL, 2014) followed by Brazil, India, Paraguay and Canada. However, it is important to point out that Brazil besides providing its internal market has a great part of its production directed to international markets. Additionally, according to USDA (20015), in the 2014/15 period, Brazil ranked as the world’s greatest exporter for the fourth consecutive year. In the last four harvests China's production became a paradox. That is, even being one of the world’s major producers was also considered the major importer.
Soybean production is considerably expanding in South America, where 50% of the world production is presently located, especially in Brazil, Argentina and Paraguay. Brazil is second in world production and maintaining the present growth tendency it may become in the coming years the first in production. In the last ten years, both Brazilian and Argentineans exports triplicated and the tendency is that Brazil becomes the major exporter (SEAB/DERAL, 2008).

**Brazilian Panorama**

In Brazil, according to Barreto (2011), soybean started to gain impulse in the 1960 years, however, it was only in the 1970 decade that production started to increase. From 1970 to 1979, soybean cultivation was more expressive in the South Region of the country and in the State of São Paulo, in view of several favorable factors for its cultivation, such as soil and climatic favorable conditions, infrastructure (roads, ports and communication), besides the foundation of an articulated soybean research net (Barreto, 2011). Even though, a significant growth in production during the 1960 years could be observed, it was only in the following decade that soybean production increased and was consolidated, from 1.5 million tons in 1970 to more than 15 million tons in 1979 as major the crop in the national agribusiness. This growth happen not only in cultivated area (from 1.3 million hectares to 8.8 million hectares), but also as an expressive increase in productivity (from 1,140 kg/ ha to 1,730 kg/ha) (Dall’Agnol; Lazarotto; Hirakuri, 2010).

By the end of the 1970 decade, more than 80% of the Brazilian soybean production still concentrated in the three states of the South Region, even though the Cerrado, in the central region of the country, gave signs that would be an important player in the soybean production process, what effectively happen starting in the 1980 decade. In 1970 less than 2% of the national production was harvested in this region and was concentrated in Mato Grosso do Sul (MS), which south part of the state presents similar climatic conditions and latitude to the north of the Paraná State. In 1980, this percent went up to 20%; in 1990, it was higher than 40% and in 2008, it contributed with 63%, tending to occupy more space on each new harvest (Dall’Agnol; Lazarotto; Hirakuri, 2010). According to Mueller and Bustamante (2002), after the 1980 decade, soybean started to gain space in other Brazilian states. Areas in the Legal Amazonia (North Region, Mato Grosso and west of Maranhão) also started to expand soybean cultivation.

By the end of 1980 and the beginning 1990 a significant increase in the national soybean production was due to its expansion in the Cerrado Region [Silva and Targino (2002)]. However, in the 1990/1991 harvest the favorable soybean grain production was interrupted by economic policy restrictions on credit affecting its production.

As years went by soybean gained even more importance in view of its growing external market demands and the increase in the Brazilian production capacity. In 2003, Brazil became the second world soybean greatest producer, representing 25% of total world production, only behind the United States. It is still occupying this position, with the exception in 2013/14, when it became the world largest producer (CONAB, 2005).

According to CONAB (2015), the beginning of soybean production in Brazil started in the seventies, with an outstanding increase in the following decades, especially after 1990. Additionally to the area expansion there was a considerable yield crop increase. In 1975, 12,145 tons of soybeans were harvested with an average yield of 1,748 kg/ha. In 2013, production reached the record volume of 86 120.8 million tons, with an average yield of 2,864 kg/ha (CONAB, 2015).

Therefore, it should also be notice the important participation of soybean production in the states of Mato Grosso, Paraná, Rio Grande do Sul, Goiás, Mato Grosso do Sul, Bahia, Minas Gerais, São Paulo, Maranhão, Tocantins, Piauí and Santa Catarina, following this order (CONAB, 2013). Soybean is one of the most important agricultural economic activities reflecting itself in pork and poultry production, regarding its weight in large extent as a significant livestock input.

Additionally, other important aspect of the soybean market is its price formation. Barros et al. (1990) say that price formation in Brazil occurs from the outside. In other words, internal soybean price is influenced by variations that take place in the external market so that there is no internal factor in the price formation.
formation of this agricultural product, which means that the country is a price taker from the external market.
However, studies about the internal market done by Mafioletti (2000, p. 3), show that the State of Paraná tends to internalize internal market price variations, and therefore they may be considered relevant in price formation. The justifications for this hypothesis, in terms of internal market, are given by three basic reasons: "(i) industrial capacity installed in the state; (ii) produced quantity; and (iii) geographic localization, with emphasis on Port of Paranaguá from where major soybean and derivatives products are exported"(Mafioletti, 2000, p. 3).

Price Transmission
An open economy exposes even more markets to international influence. Hence, price transmission can happen among different markets, however associated in such a way that one economic event that occurs in a determinant context may be due to an effect caused by another event. However, as studies of internal markets show us it is important to know that price transmission can also be caused by very interesting influence. For example, Sousa and Campos (2008) in their Brazilian soybean internal market study, concluded that Paraná and Rio Grande do Sul, Mato Grosso and Rio Grande do Sul states are co-integrated in relation to soybean prices and existence of price transmission among them in the long run. According to Saadi (2011), the explanation of the complementarily and substitution relations among different commodities can be given by price transmission. Therefore, space price transmission may occur by arbitrage directrix and in view of the Law of One Price. In this sense, other authors such as Listori and Esposti, (2012), developed space price transmission studies among different agricultural commodities. On the other hand, Hassounah, Serra and Gil (2011), analyzed petroleum prices influence on agricultural prices, and at last, price transmission of corn price future market to spot market.

Following the relevance for the research developed by price transmission in different countries of a same market, one can understand, according to Stigler (2011), the existence of commercialization policies influence. Hence, Tangermann (2011) did studies related to import and export taxes, import barriers looking into these policies among countries, especially involved in labor and as a change of these influences in price transmission and in volatility.

Using a VAR model, Block et al. (2012) analyzed in their study price transmission in the alcohol and sugar sector and concluded that sugar and sugarcane price are influenced by ethanol. However, they do not influence hydrated ethanol price but receiving influence by their own price.

Silva, Franscalori and Maia (2005), concluded in their study, also using a VAR model, that the domestic spot soybean price is explained in 32.97% by the American soybean price. What indeed demonstrates the influence in a given price transmission among open markets.

The Law of One Price
In the past literature of the Law of One Price, not many researchers were able to empirically show any evidence of such theory. However, after the publication of the subject by Isard (1977) and Richardson (1978), other studies were done apparently indicating no empirical sustainability of the Law of One Price (LOP). Later on, with the purpose to correct those mistakes and present empirical evidences of the LOP theory, research were developed using co-integration tests, which are most indicated for long run relation tests, among unitary roots variables.

Consequently, Stock and Watson (1988) concluded that co-integrated variables have the same stochastic tendency, which proportionate a better understanding of the co-integration concept. According to Engle and Granger (1987) the definition of this concept is given by the econometric method of testing the relation of two non stationary temporal series.

Years ahead, in order to test the validation of the Law of One Price, many other studies were developed in this area. The study done by Lacerda (2009) that analyzed the behavior of soybean prices in the Brazilian market vis à vis the Chinese market is a good example. The results were obtained by co-integration tests, which confirmed the validity of this theory for soybean prices among Brazil and China, known to be the largest importer of this commodity in the world.
The validity of LOP was also investigated by Lima and Burnquist (1997), when they studied the grain and flour soybean prices in the international market, using the Johansen Method between the January of 1985 and December of 1995 period. This procedure was used in a model taking into account prices of Brazil and the United States (export countries) and Germany (import country). The results confirmed that the LOP should be rejected for the soybean grain market. However for the soybean flour the results diverged. One must be aware that the Law of One Price is a theoretical concept, and that price convergence between integrated markets consists in an empirical aspect of great relevance.

**METHODOLOGY**

**Econometric Analysis**

The econometrics used to analyze soybean price transmission among major Brazilian production states was based on a Vector Regression Model (VAR), variance decomposition analysis and impulse response function analysis.

The VAR model can be expressed in the following way:

\[
X_t = A_0 + A_1 X_{t-1} + \ldots + A_p X_{t-p} + B_0 Z_t + B_1 Z_{t-1} + \ldots + B_p Z_{t-p} + e_t
\]  

(1)

Where:

- \( A_0 \) = is an \( n \times 1 \) vector of the intercept terms;
- \( A_1, \ldots, A_p \) = is an \( n \times n \) matrix of coefficients that relate the lag values of the endogenous variables to the current values of those variables;
- \( B_0, \ldots, B_p \) = is an \( n \times m \) matrix of coefficients that relate actual and lag values of exogenous variables to current values of endogenous variables;
- \( e_t \) = is an \( n \times 1 \) vector of error terms.

Each one of the \( X \) and \( Z \) variables are explained by their lag variables.

In order to choose the best VAR Model, the Schwartz Criteria (SC), Akaike Criteria (AIC) and Hannan-Quin (HQ) were used. This was done in view of the importance of the lag number determination to be included in the VAR Model, since it takes into consideration the sum square of the residues, the number of observations and the parameter's estimators. Therefore, the smaller the values the better the estimated model.

Hence, to test the stationary of the series, the ADF (Augmented Dickey – Fuller) (1979 and 1981) without structural breaking, was used in this paper to verify the integration order of the variables of interest, i.e., since it was necessary to verify the existence or not of unitary roots in the temporal series.

The next step was to test the existence of co-integration among the variables analyzed in this paper. The co-integration identifies if non stationary processes presents long run equilibrium relation, i.e., two or more non stationary time series are co-integrated if there is a long run stable relation of stationary residuals. The co-integration tests among two or more economic series allow us to accept or reject the existence of long run relations among these variables.

Before testing for co-integration, the variables integration order must be verified by using the Augmented Dickey - Fuller (ADF) unitary root test. It is necessary to observe if the series are integrated in the same order since the variables must have the same integration order.

To verify the existence of co-integration among a set of economic variables the Johansen & Juselius Method (1990) was used. This method is based on the following modify version of a VAR model:
\[ \Delta y_t = \Gamma_1 \Delta y_{t-1} + \ldots + \Gamma_{p-1} \Delta y_{t-p+1} + \Pi y_{t-1} + \varphi d_t + \mu + \varepsilon_t \]  

(2)

Where:

- \( y_t \) = a vector with \( k \) variables;
- \( d_t \) = a vector of dummy variables used to capture stationary variations;
- \( \varepsilon_t \) = the random error.

If \( r \) is the rank of the matrix \( \Pi \), then \( \Pi \) has \( r \) characteristic roots (eigenvalues) or auto values statistically different from zero. Three different situations may occur: (a) if \( r = k \), then \( y_t \) is stationary; (b) if \( r = 0 \), then \( \Delta y_t \) is stationary; (c) if \( 0 < r < k \), then \( \alpha \) and \( \beta \) matrices are such as \( \Pi = \alpha \beta' \) and the \( \beta y_t \) vector is stationary. Where \( a \) represents the velocity of adjustment of the matrix parameters in the short run, while \( \beta \) is a co-integration matrix of coefficients in the long run.

The null hypothesis of the existence of a co-integrated vector is tested by using trace statistic (\( \lambda_{\text{trace}} \)) and the maximum auto value statistic (\( \lambda_{\text{max}} \)). The trace test is given by:

\[ \lambda_{\text{trace}} = -2 \ln(Q) = -T \sum_{i=r+1}^{n} \ln(1 - \lambda_i) \]  

(3)

Where: \( Q = (\text{maximum restricted likelihood function} / \text{maximum non restricted likelihood function}) \)

The maximum auto value test is given by:

\[ \lambda_{\text{max}} = -T \ln(1 - \lambda_{r+1}) \]  

(4)

Where \( \lambda_i \) are the estimated values of the characteristic roots obtained from the \( \Pi \) estimated matrix and \( T \) is the number of observations. If the calculated values of \( \lambda_{\text{trace}} \) and \( \lambda_{\text{max}} \) are greater than the critical values, then the null hypothesis of no co-integration is rejected.

The up to here described procedures were useful to determine the long run equilibrium relationship among variables. Hence, Engle and Granger (1987) demonstrated that even presenting a long run equilibrium relationship among no stationary variables (in level), it is possible to occur a disequilibrium in the short run, i.e., the short run dynamic may be influenced by the magnitude of the deviation in relation to the long run equilibrium.

Following the co-integration test, given that the series have a common dynamic and are integrated in the same order, a more complete VAR, known as Vector Error Correction Model (VECM), can be specified since its variables with common dynamics have a long run and a short run component.

Next the Granger Causality is tested, where "the future cannot influence the past or the present" is assumed. However, if the \( Z \) variable may cause changes in the \( Y \) variable, then the \( Z \) variable precedes and/or may be relevant for \( Y \). The basic hypothesis is that the relevant information for prediction is exclusively included in the time series of these variables (Madala, 1999). Therefore, in this test two basic procedures must refer to the presence of autocorrelation among the regression residuals and the lag numbers. If autocorrelation indeed exists, it then must be eliminated. In this sense, according to Guajart (2006), a number of different filters may be used.

Along with the econometric specifications described above, the impulse-response functions to capture the intensity and the interregional price shock signal in the soybean market among the Brazilian states of Mato Grosso, Paraná, Rio Grande do Sul and Goiás were also estimated in this paper. For the unitary root
test, co-integration, VAR Model estimation, Granger causality and response-impulse the EVIEWS 8.0 was used.

The estimated model was developed in order to evaluate the hypothesis of casual likelihood among prices, received by farmers in the states of Mato Grosso, Paraná, Rio Grande do Sul e Goiás (using a soybean sac of 60 kg as unit). Time series of monthly average prices for the January 2008 to December of 2015 period were obtained from the Companhia de Nacional de Abastecimento (CONAB). This means the analysis of a period with up to 96 observations. Additionally, the series were also deflated by the IGP-DI/FGV index. The regions were chosen based on the ranking of Brazilian major crops as well as on data availability. The price series were transformed into logarithms so that the obtained beta coefficients can correspond to price transmission elasticity.

RESULTS AND DISCUSSION

The historical series of soybean price behavior in the analyzed states (Mato Grosso, Paraná, Rio Grande do Sul e Goiás) are presented on Table 1. The prices were deflated by the IGP-DI/FGV index based on December of 2015. As it can be seen, the series follow the same tendency, which may be understood that they move together in time.

First, before doing the co-integration test, it was necessary to test the stationary hypothesis or unitary root, in order to verify if the series in the natural logarithm form were stationary, that is, if they have constant mean and variance. Therefore, the Augmented Dickey-Fuller (ADF) and Phillips Perron (PP) tests were done. The results are presented in Table 1.

### Table 1: Unitary root tests: Augmented Dickey-Fuller (ADF) and Phillips Perron (PP) for soybean price series in Mato Grosso, Paraná, Rio Grande do Sul and Goiás.

<table>
<thead>
<tr>
<th>Augmented Dickey-Fuller (ADF)</th>
<th>Variable</th>
<th>Level</th>
<th>First difference</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Variable</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>( \tau_c )</td>
<td>( \tau_c )</td>
</tr>
<tr>
<td>LMT</td>
<td>-3.33**</td>
<td>-3.61**</td>
<td>-0.0672</td>
</tr>
<tr>
<td>LPR</td>
<td>-2.7***</td>
<td>-2.9624</td>
<td>-0.2050</td>
</tr>
<tr>
<td>LRS</td>
<td>-2.3251</td>
<td>-2.6277</td>
<td>-0.0793</td>
</tr>
<tr>
<td>LGO</td>
<td>-2.97**</td>
<td>-3.2***</td>
<td>-0.1924</td>
</tr>
</tbody>
</table>

Source: Based on the research results.
Variable | Level | First difference
---|---|---
| $\tau_c$ | $\tau_{ct}$ | $\tau$ | $\tau_c$ | $\tau_{ct}$ | $\tau$
---|---|---|---|---|---
LMT | -2.7*** | -2.9701 | 0.0370 | DLMT | -6.500* | -6.473* | -6.534*
LPR | -2.4122 | -2.5634 | -0.0948 | DLPR | -6.688* | -6.675* | -6.725*
LRS | -2.1727 | -2.3691 | 0.0699 | DLRS | -7.545* | -7.511* | -7.587*
LGO | -2.5*** | -2.7877 | -0.1298 | DLGO | -6.429* | -6.425* | -6.463*

**Source:** Based on the research results.

**Note:** * Significant at the probability level of 1%; ** Significant at the probability level of 5%, *** Significant at the probability level of 10%; $\tau_c$ is a statistic with a Constant, $\tau_{ct}$ is a statistic with a Constant and tendency and $\tau$ is a statistic with no Constant and tendency.

The results of the Augmented Dickey-Fuller (ADF) and Phillips Perron (PP) unitary root tests indicated that, some logarithm price series do not present specifications of the tests (with constant, with constant and tendency, and without constant and tendency) of a stationary character, when analyzed in level. However, for the first difference all series show stationary specification for any of the Augmented Dickey-Fuller (ADF) and Phillips Perron (PP) tests. Therefore, the conclusion is that the series in this study are an integrated processes of order one.

Next the soybean price co-integration analysis was done by using trace statistics and the highest auto value. Co-integration reflects a long run equilibrium in which the system converges, that is, the analyzed series presents a long run relationship if they are co-integrated indicating the existence of a vector error correction (VEC) that ties the series. Through the Akaike (AIC) and Schawarz (SC) criteria, the utilization of two lags were determined in all models, with the exception of the soybean price relation between the Rio Grande do Sul and Goiás which presented four lags. Table 2 presents the Johansen co-integration tests results for the bivariate models.

**Table 2 - Results of the Johansen co-integration test for monthly soybean price series in the states of Mato Grosso (LMT), Paraná (LPR) and Rio Grande do Sul (LRS) and Goiás (LGO), from January of 2008 to December of 2015.**

<table>
<thead>
<tr>
<th>Relation</th>
<th>Lag</th>
<th>Null Hypothesis</th>
<th>Trace Statistic</th>
<th>Highest Auto Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>LMT-LPR</td>
<td>2</td>
<td>$r = 0$</td>
<td>29.09729*</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>5.898581</td>
<td>6.634897</td>
</tr>
<tr>
<td>LMT-LRS</td>
<td>2</td>
<td>$r = 0$</td>
<td>26.11659*</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>4.152914</td>
<td>6.634897</td>
</tr>
<tr>
<td>LMT-LGO</td>
<td>2</td>
<td>$r = 0$</td>
<td>41.78109*</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>7.505650*</td>
<td>6.634897</td>
</tr>
<tr>
<td>LPR-LRS</td>
<td>2</td>
<td>$r = 0$</td>
<td>14.27764</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>2.335990</td>
<td>6.634897</td>
</tr>
<tr>
<td>LPR-LGO</td>
<td>2</td>
<td>$r = 0$</td>
<td>28.61566*</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>5.753006</td>
<td>6.634897</td>
</tr>
<tr>
<td>LRS-LGO</td>
<td>4</td>
<td>$r = 0$</td>
<td>23.87262*</td>
<td>19.93711</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$r \leq 1$</td>
<td>5.085975</td>
<td>6.634897</td>
</tr>
</tbody>
</table>

**Source:** Based on the research results.

**Note:** * indicates the null hypothesis significance at the level of 1%.
According to Table 2, among the analyzed pairs, no co-integration was verified only for the LMT x LGO and LPR x LRS relations. Therefore, the first presented a complete post and the second a null post both showing the lack of co-integration. In the remaining cases, the null hypothesis indicates that the existence of at least one co-integrated equation may not be rejected. In other words, the presence of a reduced post was verified, so that in the long run relation among these prices exists. Therefore, the results confirm the validity of the Law of One Price.

The stationary condition and bivariate co-integration among the series being defined, the VAR model was estimated. Again the Akaike (AIC), Schwarz (SBIC) and Hannan-Quinn (HQ) information criteria were used to select the optimum lag number of the VAR Model. According to the information criteria results, described in Table 3, the option was to work with only one lag.

<table>
<thead>
<tr>
<th>Lag</th>
<th>AIC</th>
<th>SC</th>
<th>HQ</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>-16.47771</td>
<td>-16.3643*</td>
<td>-16.43205</td>
</tr>
<tr>
<td>1</td>
<td>-16.80894*</td>
<td>-16.24207</td>
<td>-16.58068*</td>
</tr>
<tr>
<td>2</td>
<td>-16.66475</td>
<td>-15.64438</td>
<td>-16.25388</td>
</tr>
</tbody>
</table>

**Source:** Based on the research results.

The Vector Auto-Regression Model (VAR) was then estimated by its first difference with one lag. The results of the VAR Model are presented in Table 4.

<table>
<thead>
<tr>
<th>Coefficients</th>
<th>DLMT t-1</th>
<th>DLPR t-1</th>
<th>DLRS t-1</th>
<th>DLGO t-1</th>
</tr>
</thead>
<tbody>
<tr>
<td>DLMT t-1</td>
<td>0.455753*</td>
<td>0.281236**</td>
<td>0.202792***</td>
<td>0.274031**</td>
</tr>
<tr>
<td></td>
<td>[2.67329]</td>
<td>[2.20477]</td>
<td>[1.75409]</td>
<td>[2.11420]</td>
</tr>
<tr>
<td>DLPR t-1</td>
<td>0.675300***</td>
<td>0.471566****</td>
<td>0.860033*</td>
<td>0.726742**</td>
</tr>
<tr>
<td></td>
<td>[1.75143]</td>
<td>[1.63462]</td>
<td>[3.28927]</td>
<td>[2.47918]</td>
</tr>
<tr>
<td>DLRS t-1</td>
<td>-1.056115*</td>
<td>-0.734945*</td>
<td>-0.930943*</td>
<td>-0.703657**</td>
</tr>
<tr>
<td>DLGO t-1</td>
<td>0.185457</td>
<td>0.232201</td>
<td>0.104806</td>
<td>0.074298</td>
</tr>
<tr>
<td></td>
<td>[0.83895]</td>
<td>[1.40389]</td>
<td>[0.69914]</td>
<td>[0.44208]</td>
</tr>
<tr>
<td>C</td>
<td>0.001310</td>
<td>-0.000178</td>
<td>0.000885</td>
<td>-0.000200</td>
</tr>
<tr>
<td></td>
<td>[0.20947]</td>
<td>[-0.03813]</td>
<td>[0.20868]</td>
<td>[-0.04196]</td>
</tr>
</tbody>
</table>

**Source:** Based on the research results.

**Note:** * Significant at the probability level of 1%, ** Significant at the probability level of 5%, *** Significant at the probability level of 10%, **** Significant at the probability level of 11%.

From the VAR estimation analysis, in order to detect the influence of one variable in relation to the others, the results show that not all coefficients were statistically significant, with the exception of the soybean price in Goiás. The reason for that is that Goiás is the less significant state in terms of production among other states considered in the analysis. Hence, justifying its smaller influence. On the other hand, the price in Goiás was influenced by all other regions.

It can also be seen in Table 4, that in all regressions the soybean price relative coefficients for Mato Grosso, Paraná and Rio Grande do Sul, that lagged in one period, were significant, which means that, the soybean prices in the previous period (t-1) in those states, have influence on prices in the present period (t). This can be explained by the fact that these three states are national production reference of the commodity.

According to Figure 2 shown below, it can be noticed that the estimated model attends the stability presumption, which makes its empirical application fully confident.
The validation of the VAR stability conditions depends on the autoregressive characteristics of the polynomial inverse roots showed previously. Hence, it has been verified that all roots are inside the unitary circle, what guarantees the model stability, and therefore, the analysis of the Impulse Response function becomes possible.

With the confirmation of the stability of the estimated model the Lagrange multiplier is tested with the objective to examine the existence of autocorrelation in the VAR residuals. The results of this test are presented in Table 5.

Table 5: Lagrange Multiplier Test

<table>
<thead>
<tr>
<th>Lag</th>
<th>LM-Stat</th>
<th>Prob</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>13.46491</td>
<td>0.6385</td>
</tr>
</tbody>
</table>

Given the null hypothesis \( H_0 \) of lack of residual autocorrelation, the null hypothesis of no autocorrelation among the residuals in the first lag may not be rejected for all significant levels was verified.

After the VAR estimation, as described in Table 6, the Granger Causality test was done in order to verify the existence a casual relation among the variables.

Table 6: Granger causality test

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Independent Variable</th>
</tr>
</thead>
<tbody>
<tr>
<td>DLMT</td>
<td>DLPR</td>
</tr>
<tr>
<td>DLMT</td>
<td>-</td>
</tr>
<tr>
<td>DLPR</td>
<td>0.0275**</td>
</tr>
<tr>
<td>DLR5</td>
<td>0.0794***</td>
</tr>
<tr>
<td>DLGO</td>
<td>0.0345**</td>
</tr>
</tbody>
</table>

According to Table 6, the bidirectional causality in the Granger sense was verified among the following three relations: DLMT \( \times \) DLPR, DLMT \( \times \) DLR5 e DLPR \( \times \) DLR5. In other words, by the Granger Causality
analysis unifying the estimates previously obtained in the VAR model, the existence of price transmission among the three major producers in the country was therefore confirmed. On the other hand, with respect to the soybean price in Goiás, a unidirectional causality relation prevails, that is, there is influence of prices of other states in the prices of Goiás and not the opposite happens, showing that there is no price formation in this state, but Goiás is a price taker. Goiás has no price transmission power, but, allows price transmission from other states.

After the Granger Causality test was done, the dynamic relation among variables could be observed by the impulse response function (FARIAS, 2008). In view of the relevance of the soybean price in Mato Grosso, Paraná and Rio Grande Sul confirmed by the tests done previously, the chosen option was to obtain their respective responses to exogenous shocks. The results of the impulse response function analysis are shown in Figure 3.

![Figure 3](image)

**Figure 3**: Graphics of soybean price response of Mato Grosso in relation to price impulse of other states.

**Source**: Based on the research results.

Figure 3 presents the effects on soybean prices in Mato Gross as a result of exogenous shocks on the price of the other states. It can be seen that the price impulse in Paraná has a negative effect, almost insignificant, stabilizing itself in the fourth period. On the other hand, the Mato Grosso soybean price effect resulting from a Rio Grande do Sul price shock was negative, entering in equilibrium in the fifth period. In relation to the shock of the price in Goiás, a positive effect was verified, and after the sixth period it tends to enter in equilibrium.

![Figure 4](image)

**Figure 4**: Graphics of soybean price response of Paraná in relation to price impulse of other states.

**Source**: Based on the research results.

Figure 4, presents the effects on the Paraná soybean price as a result of price impulse in other states. A shock in the Paraná price has a positive effect of 0.04 in magnitude, becoming stable in the sixth period. The effect of the soybean price in Paraná as a result of a shock in the prices of the Rio Grande do Sul was negative, tending to the equilibrium after the fifth period. With respect to the price in Goiás, a positive effect was verified, however not very significant, tending to the equilibrium after the fourth period.

It is interesting to mention that the soybean prices in the national market are exposed to impulses that, according to Tonin and Barczcz (2008), may be affected by both the supply and demand side. On the supply side, the factors that can contribute for changes in the produced quantity are given by climatic effects variations, as well as, frost, rain lack or in excess, or pathogenic or etymologic infestation among...
others. On the demand side, economic policy instruments, such as the exchange rate, monetary policy, among others that induce modifications in the income levels, consumption habits, among others.

**Figure 5:** Graphics of soybean price response of Rio Grande do Sul in relation to price impulse of other states.

*Source:* Based on the research results.

Finally, analyzing the responses of the Rio Grande do Sul, presented in Table 5, it could be observed in the initial periods, exclusively positives effects for all shocks that were stabilized after the fifth period. The highest magnitude response verified was the one related to an exogenous shock price of Mato Grosso.

**FINAL CONSIDERATIONS**

The purpose of this paper was to exam price transmission and its causality among four major Brazilian soybean production regions, using a Vector Auto-regression Model (VAR) along the January of 2008 to the December of 2015 periods.

The results of the bivariate co-integration test indicated the existence of long run equilibrium among the following pairs of Brazilian states: Mato Grosso and Paraná, Mato Grosso and Rio Grande do Sul, Paraná and Goiás, and Rio Grande do Sul and Goiás. The existence of co-integration among these pairs is the sufficient condition in order to assure that the system converges, validating, therefore, the theoretical assumptions of the Law of One Price. The co-integration test also confirmed the existence of integration among states. In this way, for these four states, any disequilibrium in the system will by itself quickly be adjusted.

The VAR model estimates showed that soybean prices in all analyzed states were significantly influenced by passed soybean prices in Mato Grosso (MT), Paraná (PR) and Rio Grande do Sul (RS). The Granger causality analysis confirmed these results so that the MT, PR and RS prices influence the prices of all the states analyzed. Hence, there is strong evidence not to reject the hypothesis that soybean prices in MT, PR and RS are transmitted to other soybean agribusiness sectors.

With respect to the soybean price in Goiás, no expressive relations were observed with prices of other states. This can possibly be justified in view of the fact that Goiás, among the analyzed states, is the one with less representation in terms of soybean production. Then it can be assumed that: greater participation in production leads to a higher integration degree. Therefore, the State of Goiás can then be considered a soybean price taker given that it is influenced by all other analyzed states.

By the Impulse Response Function analysis, it was verified, in general, that exogenous shocks generate effects on the dynamics of soybean prices in MT, PR and RS. Effects that in the average dissipate after five months. As expected, responses of higher magnitude were caused by price shock in Mato Grosso, confirming the greater relevance of this state in soybean price transmission.

However, other factors may contribute to the strengthening of this crop, such as growing consumption diversification and higher integration among regions. Hence, other unfold researches are recommended using different variables and methodologies that could analyze space soybean price transmission.
REFERENCES

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