Article – Economic letters

Price Transmission between the International Market of Soybean Chicago (CBOT) and the Uruguayan domestic market

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Abstract: This study analyse the dynamics of price transmission between the international soybean market of the Chicago Board of Trade (CBOT) and Uruguay's domestic market, with a focus on determining the existence and nature of a long-term equilibrium relationship. Additionally, it examines the causal interactions and short-term relationships between these markets. Employing a price transmission econometric approach. The research first utilizes the Augmented Dickey-Fuller test (ADF) with structural breaks to assess stationarity within the time series. Subsequently, the Johansen Cointegration test is applied to ascertain the presence of long-term relationship among the variables. The study advances by implementing the Granger causality test to elucidate the directionality of causation and employs the Vector Error Correction Model (VECM) to dissect the short-term dynamics of the relationship. The period under study spans from January 2010 through December 2016.The analysis reveals a cointegration between the Uruguayan and Chicago markets, signifying a long-term relationship. It is observed that price alterations in Chicago impact the Uruguayan market but not vice versa, establishing a unidirectional causality. The VECM indicates that after a perturbation, the Uruguayan market adjusts towards equilibrium at a rate of 32% per month, signalling a return to equilibrium within approximately three months. These findings corroborate the Law of One Price in the long run, whereby the markets in question tend to converge to an equilibrium price. The study substantiates Uruguay's position as a "price taker," underlining its substantial reliance on international markets and confirming the efficiency of its domestic soybean market in terms of price transmission mechanisms.

Keywords: Price transmission, Uruguay, Soybean, Chicago board of trade, CBOT, Econometrics , Agricultural economics.; VECM; IRF.

Academic Editor: **Alfonso Freiria del Portillo**

ISSN (Pending)

ISSN (en trámite)

Received: 21/09/2023

Revised: 15/10/2023
Accepted: 05/11/2023
Published:

1. Introduction

Extensive research has delved into the dynamics of soybean price transmission, defined as the propagation of price changes along the supply chain or between markets. Such studies have typically scrutinized horizontal price transmission (among different markets), vertical price transmission (throughout the supply chain), and cross-price transmission (among diverse commodities), contributing to a vast body of literature on these themes. Despite the comprehensive analysis, most inquiries have focused on the predominant forces in the soybean arena: Brazil, Argentina, the United States, and China. Yet, as soybean cultivation spreads to new regions, emerging exporters like Paraguay, Uruguay, and Bolivia are entering the international stage. This shift reveals a significant void in the literature regarding these countries' soybean price transmission within their supply chains and in connection to the global market. It is crucial, therefore, to initiate research that generates insights and data for a more nuanced understanding of these dynamics.

This study centres on Uruguay, which contributes a modest 1% to global soybean production. With no substantial domestic processing industry, Uruguay exports 95% of its harvested soybeans, making its economy highly dependent on international market trends (Rodriguez, 2012). Our investigation seeks to discern the nature of price transmission and the presence of cointegration, a long-term equilibrium relationship, between the international soybean market, epitomized by the Chicago Board of Trade, and Uruguay's domestic market. We also examine the short-term interaction between these markets, assessing how price variations or market shocks in one can affect the other, and validate the application of the Law of One Price (LOP) and Uruguay's role as a "price taker" through empirical evidence.

This analysis utilizes a range of econometric models, which are expounded upon within the study. Employing quantitative secondary data, specifically monthly price series from January 2010 to December 2016, our methodology employs the Augmented Dickey-Fuller (ADF) test, with and without structural breaks to assess the integration level of the price series and to verify stationarity or the presence of a unit root.

The Johansen cointegration test was used to check the if there is a long-term relationship between the studied variables, while the Granger causality was deployed to reveal the short-term causal relationship between Chicago market and the Uruguayan domestic market. An Vector Error Correction Model was used to represent the short-term dynamics, speed of adjustment and efficiency of the market.

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1.1. Soybean; International context

The world’s most important cultivated oilseed is soybean. This valuable crop has several proposed uses, such as energy as biodiesel and human food as vegetable oil. In addition, due to the high protein content of around 38% (INTA Rafaela, 2017), soybeans are also an ideal raw material for animal feed production. Ferreira and Silva (2016) also cite other uses, for example use in the steel and pharmaceutical industries. This wide diversification of products is possible since the soybean processing industry generates two different products: oil and soymeal. These sub-products represent a valuable input for many different industries.

Soybean is classified and traded as a commodity and is produced in high volumes; production had reached 147 million metric tons per year by 2016/2017 (USDA report, 2017). The main platform through which the commodity is bought and sold is the Chicago Board of Trade (CBOT); Rotterdam Port, though much less important, represents the main entrance of the product to the European Union. In Table 1.0 illustrates the world’s balance of soybean in terms of beginning stocks, total production, imports, domestic crush, exports and ending stocks by country in two different periods of time: 2015/2016 and 2016/17.

Table 1. World Soybean Supply and Use (Million Metric tons).

World Soybean Supply and Use 1/
(Million Metric Tons)

2015/16	Beginning Stocks	Production	Imports	Domestic Crush	Domestic Total	Exports	Ending Stocks
World 2/	77.52	313.71	133.33	274.93	314.35	132.46	77.74
United States	5.19	106.86	0.64	51.34	54.47	52.86	5.35
Total Foreign	72.33	206.85	132.68	223.60	259.87	79.60	72.39
Major Exporters 3/	50.67	164.73	1.12	86.69	94.54	71.79	50.19
Argentina	31.70	56.80	0.68	43.27	47.56	9.92	31.70
Brazil	18.93	96.50	0.41	39.75	43.25	54.38	18.20
Paraguay	0.02	9.22	0.01	3.60	3.64	5.31	0.29
Major Importers 4/	18.68	15.47	113.50	106.59	128.02	0.32	19.30
China	17.01	11.79	83.23	81.00	95.00	0.11	16.91
European Union	0.68	2.32	15.12	15.20	16.83	0.14	1.15
Japan	0.21	0.24	3.19	2.28	3.38	0.00	0.26
Mexico	0.07	0.33	4.13	4.40	4.43	0.00	0.10
2016/17 Est.							
World 2/	77.74	351.25	143.61	288.40	330.28	147.46	94.86
United States	5.35	116.92	0.61	51.68	55.52	59.16	8.20
Total Foreign	72.39	234.33	143.01	236.72	274.76	88.31	86.66
Major Exporters 3/	50.19	185.78	1.97	89.33	97.49	79.51	60.95
Argentina	31.70	57.80	1.70	43.88	48.33	6.90	35.97
Brazil	18.20	114.10	0.25	41.30	44.95	63.14	24.46
Paraguay	0.29	10.67	0.01	3.95	3.99	6.60	0.37
Major Importers 4/	19.30	16.81	121.37	112.18	134.98	0.37	22.13
China	16.91	12.90	92.50	87.00	102.00	0.12	20.19
European Union	1.15	2.38	13.20	14.20	15.84	0.20	0.69
Japan	0.26	0.24	3.20	2.30	3.46	0.00	0.24
Mexico	0.10	0.51	4.20	4.65	4.69	0.00	0.12

Source: USDA, 2017

The table 1 shows that Brazil exports the most soybeans: approximately 54 million metric tons (41% of total exports) and 63 million metrics tons (42.8% of total exports) of soybeans in 2015/2016 and 2016/2017 respectively. The second main exporter is the USA with 52 million metric tons (almost 40% of total exports) in the 2015/2016 period and 59 million metric tons (40% of total exports) by 2016/2017. The next main player concerning exportation share is Argentina; this country exported roughly 10 million metric tons (around 7.5% of total exports) in 2015/2016 and 6.9 million metric tons (4.7 % of total exports) in 2016/17.

The main importer of soybean is China. In the 2015/16 period, China imported 83 million metrics tons, 73% of the world's importation of soybean, and in the 2016/2017 period, it imported 92.50 million metrics tons, representing 76% of the world's soybean importation. The second placed importer is the European Union with approximately 15 million metric tons (11.3% of the world's soybean importation) in 2015/2016 and 13.20 million metric tons (10.9%) in 2016/17.

The main players can be identified on both the supply and demand side; each player has a key role to play in the dynamics of the soybean market. There only 5 main players. On the import side, the two main importers are China and the European Union, and representing the main exporters are Brazil, Argentina and the USA (Ferreira and Silva, 2016).

Soybean is the main oilseed produced in the USA with 90% of the oilseed production (USDA 2017). The production of soybean started in the beginning of 1900 and the expansion of this crop has been constant. In 2009, soybean was the second main field crop planted only overcome by corn. The low production cost, the rising yields, the possibility of rotation with corn, and the intern demand for biofuel helped the expansion of Soybean production. In terms of commercialization USA holds The Chicago Board of Trade (CBOT), was the first and old futures market of the world. The last 10 years this market presented a high growth, currently commercialize around 50 different types of commodities including soybean. CBOT has 3 different independent markets: Spot market, Future market and Option market. The 50% of the world future contract are commercialized in this market thus CBOT is the most important commodity market in the world and so the USA market is classified in the price maker category (Silva *et al* .,, 2005).

In the 1960's, soybean arrived in Brazil at the same time as the country was experimenting with the Green Revolution. The country changed the policy of supporting family agriculture and adopted an exportation support model, paving the way for the soybean expansion throughout Brazil. In 1970 the soybean expansion was consolidated, and soybean yields also increased (Silva, *at el*., 2005). According to Brum (2004), in the eighties, the expansion of the Brazilian soybean market decreased due to growth of the risk rates concerning this activity, despite the development of the international market and the trend towards free market trading. In the following decade the domestic situation reverted, and the expansion of the market continued, led by new technologies such as direct seeding (sowing without ploughing first). As a consequence national production grew to 20 million metric tons in 1990 and 31 million tons in 1999. In the case of Argentina according to Cadenazzi (2009), soybean was first seen in Argentina during the 1970's. In 1971/1972 harvest season, soybean occupied an area of 79,800 ha. By the beginning of 1982 the soybean area in Argentina was approximately 2 million ha. In the 1986/1987 season the amount of land dedicated to soybean in Argentina overtook that of corn, the same happened in 1991/1992 with wheat, giving soybean the status of the most important crop in the country.

China has gained buying power in the internal commodities market, and by 2015 had developed a Gross Domestic Product (GDP) in purchasing power equal to $19.5 trillion (17.2% of the world's purchasing power) (Smales, 2016). This country represents the main importer of soybeans. China's consumption of imported soybean has been increasing significantly over the past decade, in part due to the implementation in 1966 of a lower tariff policy (Zhao *et al*., 2010). According to Ye and Ma (2015), the world soybean import trade is progressively concentrating in China. This country's soybean importation has the most substantial effect on the world global food market. Soybean demand is focused on supplying protein for animal production in the country, to meet the increasing meat consumption in China (Ferreira and Silva, 2016).

This strong dependency on the international soybean supply carries several risks, mainly in the domestic market; this market could affront foreign stock prices leading to such fluctuations as occurred in 2002 to 2003. These fluctuations affected all of the Chinese companies that produced oil and caused these companies to go bankrupt. As a consequence, the state adopted intervention policies such as denying the entrance of genetical modified soybean in order to stabilize the domestic market. Soybean represents China's most internationalized commodity in free markets terms. This means the commodity does not have any importation restrictions such as tariffs, quotas, taxes or others protectionism barriers (Zhao *et al*., 2010).

1.1 Price transmission empirical work

The landscape of price transmissions within South American soybean markets has been extensively studied, providing insights into the dynamic interplay between domestic and international price fluctuations. This review synthesizes key studies from the early 1990s to the mid-2010s that have shaped our understanding of price transmission mechanisms.

The pioneering work by Aguiar and Barros (1991) initiated the exploration into price transmission asymmetry between domestic and international soybean markets in São Paulo, Brazil. Utilizing the Sims Causality and Houck's Asymmetry tests, they discovered a lag of one to four months in price adjustments, with a stronger response to price increases than decreases. Subsequently, Pino and Rocha (1994) expanded the scope to include vertical transmission, demonstrating the Brazilian market's reliance on the international market, particularly the Chicago Board of Trade (CBOT).

Lima and Burnquist (1997) assessed the applicability of the LOP between Brazil, the USA, and Germany, concluding its validity for soybeans but not for soymeal. Margarido and Souza (1998) found nearly instantaneous price transmissions from the CBOT to the Brazilian market, while transmission to Paraná soybean producers was incomplete. Margarido et al. (1999) and later Machado and Margarido (2001) revealed that price variations at the Port of Rotterdam impacted domestic markets more swiftly than those from the CBOT.

Margarido, et al. (2001) applied a more extensive econometric analysis, confirming the LOP and highlighting the Brazilian market's efficiency and its dependency on international prices. In terms of supply chain integration, Giembinsky and Holland (2008), Silva*et al.,* (2005), and Ferreira and Silva (2016) all found a long-term relationship between CBOT and Brazilian prices, with Chicago leading price movements.

Costa et al. (2006) found that Brazil's commercial policies did not impede price transmission, with a long-term pass-through rate of 57% from US to Brazilian prices. Alves et al. (2008), and Coppetti et al. (2012) explored the pass-through coefficient of exchange rates on soybean exportation, finding that changes in exchange rates had incomplete pass-through effects on exporting prices, challenging the efficacy of devaluation policies for enhancing agribusiness competitiveness.

Paz et al. (2006) noted a strong cointegration between Brazilian markets and Rio Grande do Sul but a weaker one internationally. Margarido et al. (2013) and Bini et al. (2016) examined cross-price transmissions between soybean and other commodities like oil and fertilizers, acknowledging the role of biofuels and energy costs on agricultural commodity volatility.

1.1.1. Conclusion

The literature unequivocally indicates a substantial influence of international prices, particularly from CBOT and Rotterdam, on South American domestic markets, with price transmissions being significantly affected by factors like market efficiency, policy interventions, and cross-commodity dependencies. The studies collectively affirm the LOP's validity in the long run for soybeans, albeit with nuances across different regions and market segments. These findings have not only contributed to a deeper comprehension of market dynamics but also aided in policy formulation and market prediction strategies. Despite the extensive research on price transmission in the soybean global market, a significant knowledge gap remains concerning countries that are not dominant players, yet soybeans are still vital to their economies. Understanding the intricacies of market dynamics and price transmission in these nations is crucial for a comprehensive view of the soybean industry's economic impact.

2. Methodology

The research utilizes secondary quantitative data to represent international soybean prices, specifically employing the Chicago Board of Trade's soybean futures prices as a benchmark. Complementing this are domestic price data from Uruguay, expressed in US dollars per ton, which reflect the country's average export price. These domestic figures were sourced from the "Cámara Mercantil de Productos del País," a body that compiles comprehensive data on domestic transactions, including exports and domestic consumption. This information was systematically converted into a monthly format. The time series spans from January 2010 to December 2016, encompassing a total of 83 months. Subsequently, to stabilize the variance, both the international and domestic price series were transformed into their natural logarithms.

The econometric methodology employed in this research involved a series of econometrics tests to ascertain the characteristics and relationships within the time series data. Initially, the Augmented Dickey-Fuller (ADF) test was utilized to determine the order of integration of the series, identifying whether the data were stationary or required differencing to achieve stationarity. To investigate potential causal linkages and the directionality of the relationships between the time series, the Granger causality test (Granger, 1969) was applied. This test assesses whether one time series can be used to forecast another, a vital step in understanding the interplay between variables.

For discerning the presence of long-term equilibrium relationships between the variables, the Johansen cointegration test (Johansen & Juselius, 1990) was conducted. The test's critical values were drawn from Osterwald-Lenum (1992), providing a robust framework for cointegration analysis.

Finally, to articulate the connection between short-term dynamics and long-term equilibrium, an Vector Error Correction Model (VECM) was deployed. The VECM is particularly valuable as it captures the speed of adjustment from short-term disequilibrium to long-term steady state, effectively integrating the immediate dynamics with sustained trends, as described by Banerjee et al. (1993). This dual perspective enables a comprehensive understanding of both the transient and persistent changes within the studied economic parameters.

This methodology is subject to several constraints that must be taken into consideration. Firstly, it relies on secondary data, which brings inherent limitations such as data quality and completeness that could impact the study's findings. Moreover, the dataset spans only six years, which may be insufficient for detecting long-term trends and could potentially skew the interpretation of market dynamics.

Secondly, while econometric models are powerful analytical tools, they come with their own set of challenges. These include the potential for issues such as non-equivalence of cases, the necessity of making critical decisions regarding model specification, and assumptions about the nature of the data that can influence the outcomes. Furthermore, the interpretation of results is substantially bounded by the researcher's perspective and the methodological choices made during the analysis, as noted by Brady et al. (2004). Hence, the conclusions drawn from these models should be viewed with an awareness of these underlying limitations.

3. Results

After the data was converted to natural logarithm, therefore the series will be name LNChicago and LNUruguay (Logarithm-Chicago & Logarithm-Uruguay), the first step to follow was to check for stationarity, whether the time series have a Unit root or not. To ensure this was completed the Augmented Dickey-Fuller test have been done.

3.1. Unit Root tests ADF (Augmented Dickey-Fuller test)

Table 2. Unit root test (ADF) LNChicago series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-1.92	0.32
Test critical values	1%	Level	-3.51	
	5%	Level	-2.90	
	10%	Level	-2.58	

*Mackinnon (1996) one-side p-values

The Null Hypothesis (LNChicago has a unit root) cannot be rejected, because the probability levels are higher than the critical level (0.32>0.05), this means that the probability of commit Error of Type 1 (Rejected the Null Hypothesis when is true) is 32%. Additionally the estimated t-stat is higher than the provided t-stat (-1.92>-2.9)(Table 3). The Null hypothesis is accepted, LNChicago has a unit root (non-stationary). Therefore, it is necessary to transform the series to the first difference and run the test again.

Table 3. Unit root test (ADF) LUruguay series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-1.952417	0.3072
Test critical values	1%	Level	-3.511262	
	5%	Level	-2.896779	
	10%	Level	-2.585626	

*Mackinnon (1996) one-side p-values

The Null Hypothesis cannot be rejected, the probability (0.3) is higher than the critical level (probability of commit Error Type 1). Additionally the estimated t-stat is higher than the provided t-stat (-1.95>-2.9)(Table 3), therefore it can be concluded that the series has a unit root (non-stationary). It is necessary to transform the series to the first difference and perform the test again.

Table 4 Unit root test (ADF) first difference LNChicago series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-10.6	0.0001
Test critical values	1%	Level	-3.51	
	5%	Level	-2.90	
	10%	Level	-2.59	

*Mackinnon (1996) one-side p-values

If the computed t-statistic exceeds the critical value, it can be inferred that the time series is second-order integrated, denoted as I(2). However, the obtained t-statistic of -9.698546 is less than the critical threshold of -2.897223 at the 5% significance level. Additionally, the p-value is less than the conventional alpha level of 0.05 (p-value < 0.0001, as presented in Table 4). Consequently, we reject the null hypothesis of non-stationarity. Thus, the LChicago time series is confirmed to be first-order integrated, I(1), indicating it is stationary after differencing once.

Unit root test (ADF) first difference LNUruguay series

Table 5. Unit root test (ADF) first difference LNUruguay series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-10.60	0.0001
Test critical values	1%	Level	-3.51	
	5%	Level	-2.90	
	10%	Level	-2.59	

*Mackinnon (1996) one-side p-values

The computed t-statistic of -10.599 falls below the critical value of -2.897 at the 5% significance level. With a p-value of 0.0001, as indicated in Table 5, the probability of committing a Type I error is exceedingly low at 0.01%. Consequently, we reject the null hypothesis in favor of the alternative hypothesis. This leads us to conclude that the LUruguay time series is integrated of order one, I(1), indicating it does not have a unit root and is therefore stationary.

Both price time series, LChicago and LUruguay, have been found to be stationary at the first difference.

3.2. Unit Root test (ADF) with Breaks

Structural breaks within time series data can produce spurious unit root phenomena, which may in turn lead to erroneous conclusions about cointegration. It is thus essential to rigorously investigate the presence of structural breaks in the series. Upon their identification, it may be necessary to adjust the analysis or exclude these observations to ensure the integrity of the cointegration results *Table 6.* Unit root test (ADF) with breaks LChicago series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-11.14	0.01
Test critical values	1%	Level	-4.95	
	5%	Level	-4.44	
	10%	Level	-4.19	

**Vogelsang (1993) asymptotic one-sided p-values*

(Break Date: March, 2015, Break Specification: Intercept only , Break Type: Innovation outlier)

Table 7. Unit root test (ADF) with breaks LUruguay series

			t-Statistic	Prob.*
Augmented Dickey-Fuller test statistic			-11.89	0.01
Test critical values	1%	Level	-4.95	
	5%	Level	-4.44	
	10%	Level	-4.19	

**Vogelsang (1993) asymptotic one-sided p-values*

(Break Date: March, 2015, Break Type: Innovation outlier, Break Specification: Intercept only)

The null hypothesis is rejected since the calculated t-statistic (-11.89193) is significantly lower than the critical value of -4.949133 at the 5% significance level, and the associated p-value (0.01) is below the conventional threshold of 0.05. This analytic approach, s, facilitates the detection of potential structural breaks. Ignoring these breaks could undermine the validity of cointegration models in illustrating the relationship between time series. The analysis pinpointed a structural break in March 2015, corresponding with a spike in the Chicago international market prices. This price increase was attributed to concerns similar to those in 2014, where an overcommitment in U.S. soybean sales led to a domestic shortfall, prompting preemptive overbuying of soybean futures in anticipation of a second quarter deficit.

3.3. Johansen Cointegration

After identifying the structural breaks impacting the time series, and establishing that the series are integrated of order I(1) upon first differencing, we are now positioned to conduct a cointegration analysis. This will allow us to examine the existence of a long-term relationship between the time series

Table 8. Johansen Cointegration

Hypothesized No.of CE(s)	Eigenvalue	Trace statiscs	0.05 Critical value	Prob.**
None *	0.24	25.9	15.5	0.001
At most 1	0.037	3.1	3.8	0.080

Trace test indicates eqn(s) at the 0.05 level

*denotes rejection of the hypothesis at the 0.05 level

**MacKinnon-Haug-Michelis (1999) p-values

Unrestricted Cointegration Rank test (Maximum Eigenvalue)

Hypothesized No.of CE(s)	Eigenvalue	Trace statiscs	0.05 Critical value	Prob.**
None *	0.25	22.9	14.26	0.0017
At most 1	0.037	3.06	3.84	0.080

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
*denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Table 8 reveals that the Trace Statistic exceeds the critical value at the 5% significance level (Trace Statistic > Critical Value), and the probability of committing a Type I error is only 0.1%, which is below the critical threshold of 5%. This observation pertains to both the Unrestricted Cointegration Rank Test (Trace) and the Maximum Eigenvalue test. Consequently, the null hypothesis of no cointegration is rejected in favor of the alternative hypothesis, indicating the presence of at most one cointegrating relationship. This confirms the existence of a long-term association between the two variables under consideration.

These findings are consistent with previous empirical research conducted in other South American countries, which underscores the significant long-term relationships that these markets maintain with international ones. Given that these countries, including Uruguay, are often considered price takers, this is in line with the studies conducted by researchers such as Lima and Burnquist (1997), Margarido, Fernandes, and Turolla (2001), Giembinsky and Holland (2003), Costa et al. (2006), Margarido, Turolla, and Bueno (2007), and Tonin and Barczsz (2007).

3.4. Granger Causality Test

Upon establishing the cointegration between the variables, which signifies a long-term relationship, it is imperative to ascertain the directionality of causality. Specifically, the analysis must determine whether the price movements in Chicago (LNChicago) influence those in Uruguay (LNUruguay), whether Uruguayan price dynamics (LNUruguay) have an impact on Chicago's (LNChicago), or whether there exists a bidirectional causality where each variable reciprocally affects the other.

Table 9*.* Granger Causality

Null Hypothesis:	Obs	F-Statstic	Prob.
LCHICAGO does not Granger Cause LURUGUAY	83	4.1484	0.0451
LURUGUAY does not Granger Cause LCHICAGO		0.03115	0.8604

The Granger Causality Test results (referenced in Table 9) have led to the rejection of the null hypothesis that LNChicago does not Granger-cause LURUGUAY, as indicated by a p-value below the 0.05 threshold ($p = 0.0451$). This suggests there is only a 4.5% possibility of committing a Type I error, rejecting the null hypothesis when it is actually true. Conversely, the null hypothesis that LURUGUAY does not Granger-cause LCHICAGO cannot be rejected due to the high probability of a Type I error (86%, $p = 0.8601$), which exceeds the 5% critical level. Hence, the findings indicate a unidirectional causality where LCHICAGO impacts LURUGUAY but not vice versa.

This outcome aligns with established expectations and corroborates previous research indicating that domestic soybean prices in countries that are price takers do not influence the international market benchmarks such as those represented by Chicago. Given Uruguay's relatively smaller soybean market compared to major producers like Argentina and Brazil, its influence on international prices is negligible. Consequently, the inference that LNUruguay does not exert an effect on LCHICAGO stands on solid empirical ground, supported by various studies including those by Pino and Rocha (1994), Já Lima and Burnquist (1997), Margarido and Souza (1998), and others spanning from 2001 to 2007.

3.5. Vector Error Correction Model

Prior to the implementation of the Vector Error Correction Model (VECM), it is crucial to select the appropriate scenario as elucidated by Johansen and Juselius (1990), Juselius (1990), and Johansen (1995). This decision is pivotal because the deterministic components may vary when transitioning from a Vector Autoregressive (VAR) model to a Vector Error Correction Model (VECM). Specifically, deterministic terms that are part of the cointegration relationship within the VAR framework may not be integrated into the VECM. For the purposes of this analysis, the scenario selected involves a linear deterministic trend, which must be carefully considered to ensure accurate specification of the VECM.

<u>The result the long run equation:</u>

Equation 1.

$$LURUGUAY = 0.34 + 1.05LNCHICAGO \quad (1)$$

And The vector error correction model:

Equation 2.

$$D(LNUruguay) = -0.323(1.05LNCHICAGO + 0.34) + 0.39(LNUruguay_{(t-1)}) + 0.218D(LNChicago_{(t-1)}) + 0.001921 \quad (2)$$

The error correction term's negative and statistically significant value indicates that any deviation from the long-run equilibrium caused by shocks to the independent variable (LNChicago) is corrected at a rate of 32% per month in the dependent variable (LURUGUAY). This implies that the adjustment from Chicago to Uruguay prices occurs relatively swiftly, with equilibrium typically being restored within approximately three months.

When compared with similar research conducted in other countries, this speed of adjustment is faster than the 26.16% observed in Brazil and slightly exceeds the 31.27% found in Argentina. These disparities underscore the unique market characteristics of each nation. Brazil, which boasts a substantial domestic market for soybeans as the world's leading chicken exporter, would naturally exhibit a slower reversion to equilibrium due to its robust internal demand, where soybeans are a key component of poultry feed. Conversely, Argentina's market adjustment is more prompt, reflecting a lesser domestic influence.

Uruguay's market demonstrates an even quicker adjustment than Argentina's, likely attributable to its relatively small domestic market. This expedited adjustment pace is consistent with expectations and aligns with the findings from prior research on other South American countries, reinforcing the idea that smaller markets may experience more rapid price corrections following international market shocks.

4. Conclusion

The objective of this study was to examine the price transmission between the international soybean market, as exemplified by the Chicago Board of Trade (CBOT), and the domestic market of Uruguay, utilizing econometric models and time series analysis of prices. Grounded in the Law of One Price (LOP) and the theoretical framework developed by Mundlack and Larson (1992), the investigation tested the proposition that markets, despite short-term fluctuations, converge to a uniform price in the long term. The analysis confirmed that the Law of One Price holds true for the correlation between the domestic market of Uruguay and the CBOT.

The Johansen cointegration test revealed a long-term relationship between Uruguay's domestic prices and the international prices set by the CBOT. The Granger causality analysis indicated a one-way causal effect, where price movements in Chicago have a bearing on Uruguayan domestic prices, but changes in the latter do not influence the CBOT. The Vector Error Correction Model (VECM) identified short-term deviations from the equilibrium price, demonstrating that Uruguayan prices adjust at a rate of approximately 32% monthly in response to price shocks from Chicago, returning to equilibrium within an estimated period of three months. This rate of adjustment is considerably rapid when compared to similar metrics for other countries such as Argentina, Brazil, and the USA.

The swift adjustment in Uruguay can be attributed to its relatively small and highly efficient domestic market, which contrasts starkly with countries that have a significant domestic consumption of soybeans. In Uruguay, almost 98% of produced soybean is exported, and there is minimal to no domestic soybean processing industry. This setup positions Uruguay as a highly responsive market to international price signals, underscoring its status as a price-taker within the global soybean market. The study successfully verified Uruguay's role as a price-taker in the international soybean market, consistent with previous empirical findings in other South American contexts.

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