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COINTEGRATION AND PRICE LINKAGES IN THE MERCOSUR BEEF CATTLE MARKETS

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Abstract

This research provides empirical evidence of the degree of integration of Mercosur region beef cattle markets with the international market. Prices for six spatially distinct cattle markets located in the four countries were analyzed using a fractional cointegration approach. The analysis included the computation of fractional integration parameters and the error term of the cointegration equations. The null hypothesis of 'separate markets' could not be rejected in any of the cases. This conclusion is absolute for Argentina; there is strong evidence of separate markets with Uruguay, which was chosen as a proxy of the international market. For Brazil and Paraguay, empirical results may suggest a weak degree of market integration but not enough to reject the null hypothesis. Despite the common price and industry trends in the different markets, their responses to specific price shocks were dissimilar and, most importantly, the reversion of the prices to equilibrium was always slow.
Keywords: *Cattle prices, fractional cointegration, market integration, the law of one price, economic bloc*
JEL Codes: *C5, F15, Q1*

1. Introduction

In an integrated agricultural market, a commodity's price will be determined by the forces of supply and demand, and reflect value throughout the integrated region or at the global level when a good is traded in the world market.

Under the Law of One Price (LOP), a commodity's true value (after exchange rate adjustments and accounting for transportation costs) will prevail across geographically separated markets located in one or more countries. Applying the LOP, the geographic extent of a market and the degree of integration within a market can be assessed and points of chronic market integration failure can be identified and addressed. Short-run deviations from LOP will always exist and can be explained by factors such as exchange rate volatility and other "overshooting effects" (Ardeni, 1989). Markets are dynamic and continuously adjusting; however, when allowed to function, a singular level of value will emerge and serve as a price signal throughout the integrated region. Systemic price variability throughout a region will reflect more than spatial or transitory currency-related variability when markets are not integrated.

The importance of market integration has been documented in numerous situations. Goodwin & Schroeder (1991) noted that spatial market integration in the US beef sector paralleled increased concentration in cattle slaughtering. Cattle markets can be viewed as highly differentiated markets, even within the same geographic location, due to differences in the endowment of animal traits (Buccola & Jessee, 1979; Buccola, 1980).

Market integration is central to regional and global economic development and is a principal objective of both regional trade blocs and multilateral trade agreements. The economic benefits of market integration have been touted as a key rationale for trade liberalization and significant efforts have been made globally in attempts to reduce barriers to trade, harmonize policies and regulations, and facilitate trade in goods and services. However, market integration and, more specifically, integration to international markets remain a theoretical objective in many industries and regions, and a condition that is frequently derailed by larger forces. At best, market integration should be viewed as a moving target.

In the cattle industry, spatial price relationships have important implications in defining geographic markets, promoting price discovery, and assessing market performance. A nation's beef cattle sector is dependent upon regional or national resource endowments and the primary traded good is a live animal through most of the production process. Thus, the beef cattle market and supply chain are particularly vulnerable to a wide range of market disrupters. Price shocks arise from climatic events, livestock as well as zoonotic disease threats, public perceptions, and both macroeconomic and sectoral policies. Despite many factors working against cattle market integration (international as well as regional and intra-national), there has been a gradual trend toward increased integration globally.

The *Mercado Común del Sur* (Mercosur, Southern Common Market) nations of Brazil, Argentina, Paraguay, and Uruguay account for 27% of the world's total beef herd (USDA-FAS, 2016). There is almost 290 million head of cattle in the region with Brazil accounting for more than 70% of the total (by comparison, the US cattle herd consists of ~90 million head). Argentina and Uruguay are well established as grass-fed beef exporters, and Paraguay's beef sector has expanded in recent years.

The four Mercosur countries produced 13.2 million metric tons (MMT) of beef (carcass weight) in 2013 (USDA-FAS, 2016), generating 22% of world beef production. Mercosur inhabitants lead the world in per capita beef consumption (Romero, 2013); however, the relatively small population of the region means that it will continue to be a major supplier to the rest of the world. Brazil, Uruguay, Argentina, and Paraguay all rank in the list of top beef exporters (USDA-FAS, 2016); in 2013, the bloc exported 1.8 MMT shipped weight (equivalent to 2.7 MMT carcass weight) to 120 different markets, for a total FOB value of US$ 8.8 billion. This volume represented near 30% of world's beef exports.

Table 1. Herd Size, Beef Production, and Exports of Mercosur Countries, 2013.

Country	Herd Size (million head)	Beef Production		Beef Exports			Markets
		TMT(1)	%	TMT(1)	%	Million US$	
Brazil	211.8	9,307	70.6	1,849	68.5	5,359	98
Argentina	51.0	2,850	21.6	186	6.9	1,293	59
Paraguay	13.4	510	3.8	326	12.1	991	38
Uruguay	11.5	525	4.0	340	12.5	1,177	58
Total	**287.7**	**13,192**	**100.0**	**2,701**	**100.0**	**8,820**	**120**

Source: Herd size and production based on FAOSTAT (2016) and USDA-FAS (2016). Beef exports and destination markets based on USDA-FAS (2016) and official Custom's data (URUMOL 2016). All data correspond to 2013.

Note: (1) TMT: thousand metric tons, carcass weight

Integration of the Mercosur beef market has important implications for trade harmonization, government regulation, and general economic policy (Fossati, Lorenzo, & Rodríguez, 2007). Mercosur officials issue frequent pronouncements regarding livestock market integration in the region. Overall, the beef-cattle sector in the Mercosur bloc can be characterized as land-extensive and grass fed. Despite similarities in the four countries' cattle production systems, there are important agro-ecological differences within the region (Duran, 2014).

In that sense, questions remain as to whether or not this vast cattle production region is actually a single, albeit spatially segmented, integrated market, or whether the region is comprised of separate independently functioning markets. If sub-regions are relatively independent, to what degree is the Mercosur cattle market integrated? Or, as it could be more relevant, to what *degree* are the different cattle markets of the Mercosur bloc integrated to the international market?

Uruguay may be regarded as representative of the international beef and cattle markets without too much effort, because of its high dependence on beef exports and its access to most of the major import destinations in the world. Not surprisingly, some important international consultancy firms use Uruguay's beef export price index as a "proxy" of world market conditions, in their public and private reports (GIRA, 2017).

Thus, comparing the degree of integration of each regional cattle market of the MERCOSUR region against Uruguay could be a way of assessing the degree of integration between these markets and the world market. Some other characteristics of Uruguay's beef and cattle market, which reinforce this idea, are discussed later in this article.

The Law of One Price provides a theoretical framework for measuring the degree of integration of Mercosur cattle markets. As noted by Ravallion (1986), measuring market integration provides critical information for understanding how specific markets work. As indicated above, both inter- and intra-governmental policies designed to facilitate and foment trade benefit from measurement of market integration. At a minimum, measurement provides a benchmark for assessing progress toward integration over time.

2. Market Integration and the Law of One Price

Spatial market integration concepts were developed by Takayama & Judge (1971), who provided a theoretical model based on the assumption that when information and goods flow freely, prices of a homogeneous good in two spatially separated markets should only differ by transaction costs. If the price in one market is greater than the price in a second market plus the transaction costs required to move the product from the low-price market to the high-price market, unexploited pure profits will exist. Rational traders would be attracted to the profits, enter the market, and capitalize on these arbitrage opportunities, thus reducing supply in the low-price market and increasing supply in the high price location. These two forces will, *ceteris paribus*, drive up the price in the initially low-price market and reduce the price in the market with the original higher price. Ultimately, prices adjust to the point where the price differential between the two markets equals only the transaction costs and trade between the regions only results in zero expected profits from arbitrage.

When two markets completely integrate into a single market, price changes in the exporting region induce price changes in the importing region, in the same direction and magnitude. The extent and the speed at which shocks are passed through, and the strength of the interdependence among prices are indicators of the degree of integration and the efficiency of the two markets' performance.

Cointegration analysis measures the degree of market integration between spatially distinct markets. In economic terms, two variables are cointegrated if there exists an equilibrium or a long run relationship between them (Gujarati, 2003). The appeal of using these analytical procedures to test for LOP derives from the fact that most price series are non-stationary. In other words, "they grow over time and so they do not have a fixed 'stationary' mean" (Kennedy, 1998, p. 263). Cointegration theory allows for testing long-run relationships between or among economic variables in the presence of non-stationarity.

Goodwin & Schroeder (1991) highlighted the rationale behind the use of cointegration analysis for evaluating spatial price linkages in regional cattle markets, by stating: "Economic forces should prohibit persistent long-run deviations from equilibrium conditions, although significant short-run deviations may be observed" (p.453). While noting that cointegration is not an absolute measure but a matter of degree, these authors drew attention to cointegration tests as an especially suitable framework for analyzing long-run price relationships among regional cattle markets. Weekly prices in regional cattle markets are highly variable and often possess significant trends, suggesting the potential for non-stationary behavior in long-run prices series. However, these authors argued, "efficient arbitrage and basing-point pricing conditions in regional markets suggest that regional prices in alternative markets should not diverge from one another" (p.453).

Goodwin & Schroeder (1991) enumerated several factors affecting cointegration among regional fed cattle markets: a) agent's cost and risk associated with trade between markets; b) amount of market information reflected in prices at a particular market; c) market volume; and d) packer concentration.

This integration of Mercosur beef market assesses LOP by using the fractional cointegration approach proposed by Marinucci & Robinson (2001). Fractionally cointegrated variables show more significant short-run persistence to shocks than fully cointegrated variables. The fractional cointegration analysis allows the equilibrium errors to follow a fractionally cointegrated process so that the order of integration is a fraction between zero and one.

By avoiding the discrete hypothesis of unit-roots/no-unit-roots in equilibrium, this method permits the analysis of a wider range of mean-reversion behavior than standard cointegration analysis. The long memory processes, such as the autoregressive integrated moving average models (ARIMA), have the property that their autocorrelation function decays at a much slower rate than that of a linear autoregressive-moving average (ARMA) process (Castaño, Gómez, & Gallón, 2008).

When building a model, the usual practice of differentiating the series to achieve stationarity may have negative consequences (Granger & Joyeux, 1980). While the alternative of not differentiating is also not appropriate because it would imply the non-stationarity of the series, the differentiation would generate an over-differentiation (Pérez, 2001).

To solve this problem, Granger & Joyeux (1980), Granger (1980), and Hosking (1981) introduced the so-called ARFIMA models (autoregressive fractionally integrated moving average) for modeling economic series. These models cover the "intermediate case" that exists between the unitary root ARIMA processes and the ARMA processes (Pérez, 2001).

3. Empirical Approach

3.1 Estimation Procedure

LetP_t^i denote the price of a specific good in country *i*, $P_t^j$is the price of the same good at country *j*. All the prices are expressed in a common currency using the corresponding the exchange rate, all at time *t*. Arbitrage is reached when $P_t^i = P_t^j$.

In order to simplify notation, lets rename $y_t = P_t^i$ and $x_t = P_t^j$. To allow for deviations from the LOP assumption, as well as for effects not included in the model, a disturbance term denoted by e_t can be added in order to estimate the following expression:

$y_t = \beta.x_t + e_t$. (1)

In this equation, the coefficient β does not have the same interpretation as with traditional linear regression. By construction, β is computed by using the price series (periodogram) in the frequency domain (Marinucci & Robinson,2001). Under this approach, initial differences in levels of the variables do not affect the estimation of β. Asymmetric variations in external factors, such as transportation costs, should be captured by the error term.

According to Engle & Granger (1987), two series y_t and x_t are cointegrated if a linear combination of the series has a lower level of integration. If the price series are each integrated of some order d, which is denoted as $I(d)$, yet $e_t = \beta.x_t - y_t$ is $I(d_e)$, and $d_e < d$.

Granger & Joyeux (1980) showed how fractionally differenced variables could have long-memory property. The ARFIMA processes produce long memory if the differencing parameter is in the range $0 < d < 1/2$, in which case the process is stationary and invertible. For integrated processes (*i.e.*, $d = 1$) the effect of a shock persists indefinitely while in a fractionally integrated process with $0 < d < 1/2$, the effect of a shock disappears and the series finally reverts to its mean. If $-1/2 < d < 0$, the process is said to be anti-persistent or has short-term memory because the spectral density is canceled in the origin and is dominated by the high frequencies, while the autocorrelations are all negative and absolutely addable. If $d = 0$, it will be a short-term memory process. On the contrary, if $1/2 < d < 1$, the behavior of non-stationary series that eventually revert to the mean can be modeled. This is something that unitary root processes cannot do because the process will be non-stationary.

The fractional cointegration analysis was carried out in three steps. The authors used R software environment in order to program their own code or adapt code written for previous studies (Rostán, 2009) for important parts of the estimation procedures, as the required algorithms were not available as canned routines in SAS or other common stats software.

First, the order of integration of each series, d_y, and d_x, was estimated using the semi-parametric method proposed by Geweke & Porter-Hudak (1983), known as GPH. In the case of the $\{y_t\}$ series, the problem to be solved is the estimation of parameter d_y in the following expression:

$(1 - B)^{dy}\, y_t = \mu_t$. (2)

B is the backward shift operator, so that equation (2) means that y_t is differenced at order d_y. The problem is analogous for the $\{x_t\}$ series.

Applying the GPH procedure implies the use of “frequency domain analysis” or “spectral analysis”. This means the use of spectral representations of the price series (Brockwell & Davis, 1996). Essentially, the spectral representation of a stationary time series is a decomposition of the series into a sum of sinusoidal components of a certain bandwidth, with uncorrelated random coefficients.

An important issue in the estimation of the long memory parameter d is the choice of the window width or “bandwidth”, denoted as m. Various approaches have been undertaken to overcome this problem although there is no consensus regarding the best method (Geweke & Porter-Hudak, 1983; Hurvich & Beltrao, 1994; Henry & Robinson, 1996; Delgado & Robinson, 1996; Giraitis, Robinson, & Samarov, 1997; Hurvich & Deo, 1998). All the approaches have strengths and weaknesses. A more traditional empirical criteria of setting different bandwidth values [m = 20; m = 30; m = 40] was chosen for this research.

Here, the differencing parameters d are not integers but real numbers. Comparing the point estimates of d_y and d_x, along with the comparison of their corresponding confidence intervals, makes possible testing if both price series exhibit equal order of integration.

The second step of the analysis involves the use of a regression method to estimate the fractional cointegration parameter β in the representation (1), following Marinucci & Robinson (2001). Assuming that β is identified, the e_t process is $I(d_e)$ with $d_e < d_y$. Values of β close to 1 imply that a variation in prices is fully transmitted to the domestic prices, whereas a value of $\beta = 0$ implies no transmission at all.

The third and final step consists in estimating the error term e_t and the order of integration $I(d_e)$ using the Engle & Granger (1987) approach. The order of integration d_e of the error should be lower than the order of integration of each price series. If the equilibrium error is stationary, external shocks can have a short-term impact, but little long-term effects, as the data revert to the mean of the series at an exponential rate.

The application of a long memory model, associated with a fractional integration approach, allows flexible modeling of low-frequency behavior, with the important implication of the measurement of shock persistence in the error term. The point estimates d_e are quite striking as all are less than unity and some are close to 0.5. We will discuss in the next sections the results and implications.

3.2 Data

Uruguay was designated as the baseline or "world-market" country in this cointegration analysis of the four-nation Mercosur beef market; Brazil, Argentina, and Paraguay are the "individual-market" countries. Thus, all comparison results derived from the cointegration analysis are made against Uruguay. There are some reasons supporting this choice.

First, Uruguay is the most highly beef export-dependent of the four members of the economic bloc. About two-thirds of the national production is exported every year and, in monetary terms, beef exports average something more than 17% of total exports of goods for Uruguay. On average, Paraguay exports 57% of its total beef production volume, which represent almost 15% of total exports. In turn, the volume of beef exports accounts for 17% of the national production for Brazil and 8% for Argentina. For these two countries, the value of beef exports represents no more than 2% of total national exports of goods (USDA-FAS, 2016; URUMOL, 2016).

Second, the National Meet Institute of Uruguay publishes a weekly index relating live cattle prices and the average value obtained by Uruguay beef exports (INAC, 2016). The evolution of the so-called RHE index (in Spanish, Relación Hacienda/Exportación) in the last 10 years (2005-2015) suggests that cattle prices in the local market adequately reflect world beef market conditions (GIRA, 2017).

Third, Uruguay's beef and cattle are quoted and traded in US dollars, no matter the final destination (domestic or export market). Because Uruguayan beef exporters, and cattle producers as well, trade and receive payments in US dollars, setting Uruguay (UY) as the "world-market country" in the cointegration analysis has the advantage of using the US dollar as the baseline currency and avoiding the arbitrage of two currencies with a third one. Thus, the only relevant exchange rates are the individual rates between the US dollar and the currencies of Brazil (Real), Argentina (Peso Argentino), and Paraguay (Guaraní).

Finally, Uruguay has a unique sanitary status in Mercosur. While all four countries are officially recognized as having negligible Bovine Spongiform Encephalopathy (BSE) risk by the World Organization for Animal Health (OIE according to the Spanish acronym), Uruguay is the only one recognized with the status of Foot and Mouth Disease (FMD) free with vaccination by the same international institution (OIE, 2016). Brazil, Argentina, and Paraguay, each have one or more zones declared FMD free with vaccination but do not have this status nationwide.

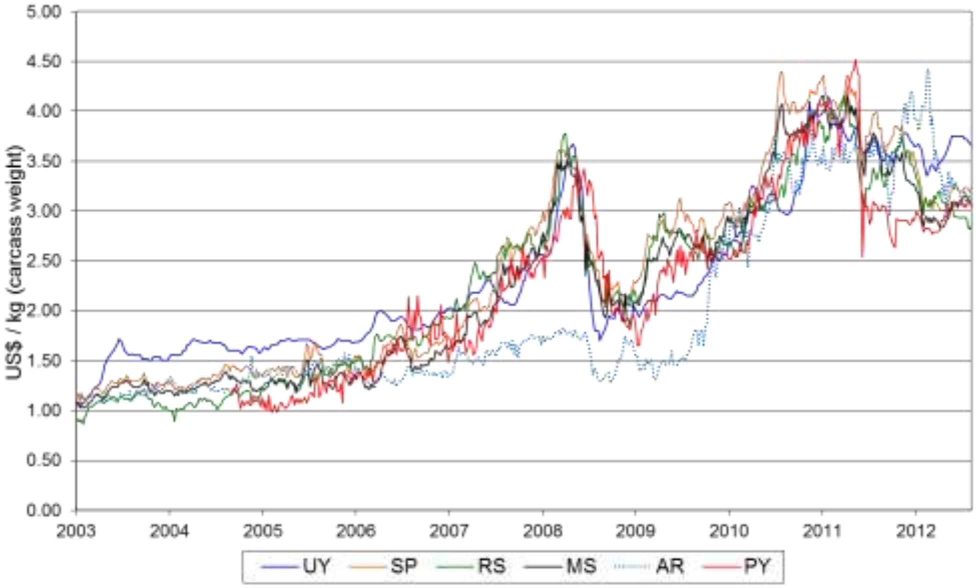

Figure1. Weekly Prices of Fat Steers: UY, SP, RS, MS, AR, and PY (2003/Apr-2012/11)

The heterogeneity of Brazil's cattle sector necessitated treating that country as separate markets rather than a single market. The three leading cattle-producing states of São Paulo, Mato Grosso do Sul, and Rio Grande do Sul are all located near the Argentine, Uruguayan, and Paraguayan markets and account for 22% of Brazil's cattle herd and 40% of that nation's beef exports.

Price series used in the cointegration analysis were obtained from local sources in each country. Uruguay (UY) prices were provided by the Instituto Nacional de Carnes (INAC); prices for São Paulo (SP), Mato Grosso do Sul (MS), and Rio Grande do Sul (RS) cattle markets were provided by the Centro de Estudos Avançados em Economia Aplicada (CEPEA) of the Escola Superior de Agricultura "Luiz de Queiroz", of the Universidade de São Paulo (ESALQ-USP) in Piracicaba, Brazil. Prices for Argentina (AR) are for the Liniers cattle market (Mercado de Liniers S.A.) and the price series for Paraguay (PY) was obtained from a private database provider.

As pointed out by Rostán (2009), precise computation of each periodogram and cross-periodogram is only possible when using data series containing at least 300 observations. This analysis was conducted using average weekly prices for fat steers ready for slaughter, quoted in US dollars/kilogram, carcass weight, spanning a period of 501 weeks, beginning the last week of April 2003 and ending the last week of November 2012. In the case of Paraguay, the available dataset covered a shorter period, from the last week of November 2004 through the final week of December 2012. Thus, the Uruguay-Paraguay cointegration analysis used only 422 observations.

4. Results and Discussion

4.1 Rule of Thumb

Robinson & Marinucci (1998) warned that there not exist well-developed inference rules for fractional stationarity and non-stationary processes. The lack of a formal standardized test for market integration using a fractional cointegration analysis requires following an *ad hoc* procedure and setting an arbitrary rule of thumb. The pairwise analysis examines the world market against each individual market with the same procedure followed for each pair of prices (*e.g.*, UY-SP, UY-RS, UY-MS, UY-AR, and UY-PY). The null hypothesis (Ho) is that the markets are separate or that the home and foreign markets are not cointegrated. The alternative hypothesis (Ha) holds that both markets are cointegrated.

The statistical analysis was carried out for three different bandwidth values of m (20, 30 and 40), to improve the level of security for interpreting the empirical results. The first condition in the empirical rule-of-considered a pair of prices as cointegrated if the result is positive for at least two values of m. Performing the analysis, pairwise, for each combination of series (UY against the others) results in a vector of economic variables $\mathbf{x}_t$. This vector is univariate and generates a scalar (x_t), as well as the value of the parameter β. The magnitude of βoffers a measure of the cointegration between the series. In addition, the reported results also include the estimated orders of fractional differencing of the series in levels (d_y and d_x) and the error of the cointegration relationship (d_e)along with the 90% confidence interval (CI) and the computed probability value (p-value).

The next step in the analysis is to verify the existence of a balance between the price series (home against foreign). In this article, each pair of prices exhibited such a balance if some range of the estimated CI of d_y is included within the estimated CI of d_x and vice-versa. If that is the case, d_y is not statistically different from d_x, the series are balanced, and the analysis proceeds to the next step. If the existence of balance is rejected, the level of integration of both series is different and they cannot be fractionally integrated.

Last, the differencing parameter of the error term (d_e) is compared with the integration orders of the series under analysis. Following Engle & Granger (1987), the value of d_e has to be smaller than the value of d_y and d_x. In addition, the behavior of the error term in the cointegration relationship is analyzed. For an unambiguous conclusion that the series are fractionally cointegrated, the error term must be stationary (d_e< 0.5); if the error (d_e) is between 0.5 and 0.75, it indicates that although it reverts to its mean (*i.e.*, equilibrium)doubts exist about the cointegration relationship because the velocity of this adjustment is very slow (Granger & Joyeux, 1980).

4.2 Empirical Results

The results of the UY-SP analysis appear in table 2. The estimated value of $\hat{\beta}$ was 0.663. In the case of m = 20, the estimated cointegration was given by d_y = 0.774, d_x = 1.040, and d_e = 0.485. According to this result, there exists fractional cointegration. The orders of differentiation of the fractional series taken in levels were not significantly different between them, as indicated by corresponding CI overlap; the degree of cointegration for the error term was less than 0.5 and lower in magnitude relative to the orders of the series in their levels.

For m = 30, d_y = 0.971, d_x = 1.020, and d_e = 0.921, while for$m$ = 40, the estimated integration orders were d_y = 1.033, d_x = 1.075, and d_e = 0.876. In both cases, the error is nonstationary (d_e > 0.75), suggesting no cointegration between the price series. In accordance with the *ad hoc* rule of thumb defined in this research (*e.g.*, at least two values of m must be positive for the series to be considered cointegrated), there is not enough evidence to reject the null hypothesis. Thus, cointegration between UY and SP cattle prices could not be established.

Table 2. Cointegration Analysis between Uruguay (UY) and São Paulo (SP) Cattle Prices

Value of *m*	Order of Integration			Cointegration Coefficient
	UY price	SP price	Error	
m = 20	d_y = 0.774	d_x = 1.040	d_e = 0.485	β = 0.663
	CI: (0.582; 0.967)	CI: (0.889; 1.190)	CI: (0.342; 0.629)	
	p-value: 0.0008	*p*-value: 0.0000	*p*-value: 0.0034	
Yes. Evidence of fractional cointegration ($d_y \cong d_x$; $d_e < 0.5$; $d_y > d_e$ and $d_x > d_e$).				
m = 30	d_y = 0.971	d_x = 1.020	d_e = 0.921	β = 0.663
	CI: (0.851; 1.090)	CI: (0.931; 1.108)	CI: (0.805; 1.038)	
	p-value: 0.0000	*p*-value: 0.0000	*p*-value: 0.0000	
No. The error is non stationary ($d_e > 0.75$)				
m = 40	d_y = 1.033	d_x = 1.075	d_e = 0.876	β = 0.663
	CI: (0.968; 1.098)	CI: (1.151; 1.151)	CI: (0.791; 0.961)	
	p-value: 0.0000	*p*-value: 0.0000	*p*-value: 0.0000	
No. The error is non stationary ($d_e > 0.75$)				

Note: 501 observations

Table 3. Cointegration Analysis between Uruguay (UY) and Rio Grande do Sul (RS) Cattle Prices

Value of *m*	Order of Integration			Cointegration Coefficient
	UY price	RS price	Error	
m = 20	d_y = 0.774	d_x = 0.960	d_e = 0.533	β = 0.607
	CI: (0.582; 0.967)	CI: (0.820; 1.101)	CI: (0.330; 0.735)	
	p-value: 0.0008	*p*-value: 0.0000	*p*-value: 0.0169	
Inconclusive. Slow adjustment to equilibrium aftershocks ($0.5 < d_e < 0.75$).				
m = 30	d_y = 0.971	d_x = 1.03961	d_e = 0.908	β = 0. 607
	CI: (0.851; 1.090)	CI: (0.931; 1.108)	CI: (0.771; 1.045)	
	p-value: 0.0000	*p*-value: 0.0000	*p*-value: 0.0000	
No. The error is non stationary ($d_e > 0.75$)				
m = 40	d_y = 1.033	d_x = 1.067	d_e = 0.904	β = 0. 607
	CI: (0.968; 1.098)	CI: (0.999; 1.135)	CI: (0.789; 1.019)	
	p-value: 0.0000	*p*-value: 0.0000	*p*-value: 0.0000	
No. The error is non stationary ($d_e > 0.75$)				

Note: 501 observations

The analysis for UY-RS appears in table 3. The estimated value for $\hat{\beta}$was 0.607. For m = 20, the integration order was given by d_y = 0.774, d_x = 0.960, and d_e = 0.533. The error integration order was greater than 0.5 but smaller than 0.75. Thus, given a shock affecting the series, they revert to their equilibrium value very slowly. For$m$ = 30, d_y = 0.971, d_x = 1.040, and d_e = 0.908; for m = 40, d_y = 1.033, d_x = 1.067, and d_e = 0.904. These results suggest that the series are not cointegrated; in addition, the error is non stationary (d_e > 0.75). According to the rule of thumb, the statistical evidence is incompatible with cointegration between Uruguayan and Rio Grande do Sul cattle prices.

A similar outcome occurred for the UY-MS comparison (table 4), since the $\hat{\beta}$ coefficient was 0.636. For m = 20, the integration order was given by d_y = 0.774, d_x = 1.049, and d_e = 0.526. Again, the error integration order was greater than 0.5 but smaller than 0.75, indicating that mean reversion after a shock is very slow.

For m = 30, d_y = 0.971, d_x = 0.981, and d_e = 0.912, while for m = 40, the integration orders were d_y = 1.033, d_x = 1.047, and d_e = 0.902. With these results, there is no evidence to reject the null hypothesis of separate cattle markets in Uruguay and Mato Grosso do Sul.

The results for UY-AR (table 5) were relatively stronger in terms of accepting the null hypothesis. In this case, $\hat{\beta}$ of 0.448 showed the lowest magnitude for the coefficient of fractional cointegration. For m = 20, the corresponding orders were d_y = 0.774, d_x = 1.173, and d_e = 0.550, but the CI estimate of d_y is not included within the CI of d_x. With these results, the series have different levels of integration order. For m = 30, d_y = 0.971, d_x = 1.089, and d_e = 0.889, revealing a non-stationary error. The same result was found for m = 40, where d_y = 1.033, d_x = 1.045, and d_e = 0.950. Following the rule of thumb, no evidence suggests the integration of Uruguay and Argentina cattle markets.

Table 4. Cointegration Analysis between Uruguay (UY) and Mato Grosso do Sul (MS) Cattle Prices

Value of *m*	**Order of Integration**			**Cointegration Coefficient**
	UY price	**MS price**	**Error**	
m = 20	d_y = 0.774	d_x = 1.050	d_e = 0.526	β = 0.636
	CI: (0.582; 0.967)	CI: (0.873; 1.226)	CI: (0.390; 0.661)	
	p-value: 0.0008	p-value: 0.0000	p-value: 0.0011	
Inconclusive. Slow adjustment to equilibrium aftershocks (0.5 < d_e < 0.75).				
m = 30	d_y = 0.971	d_x = 0.981	d_e = 0.912	β = 0. 636
	CI: (0.851; 1.090)	CI: (0.889; 1.073)	CI: (0.807; 1.018)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
No. The error is nonstationary (d_e> 0.75).				
m = 40	d_y = 1.033	d_x = 1.047	d_e = 0.902	β = 0. 636
	CI: (0.968; 1.098)	CI: (0.991; 1.123)	CI: (0.815; 0.989)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
No. The error is nonstationary (d_e> 0.75).				

Note: 501 observations

The magnitude of the estimated cointegration coefficient for UY-PY (table 6) was the highest among all the comparisons ($\hat{\beta}$= 0.967), indicating that both series follow similar pathways, except at the end of the period due to Paraguay's FMD outbreak. As noted above, the UY-PY analysis used 422 observations (*i.e.*, 79 fewer weeks than the other pairwise comparisons). Thus, the computed d_y for UY prices were different for all values of m. With m = 20, the integration orders were d_y = 1.043, d_x = 1.146, and d_e = 0.69, with the result showing slow post-shock adjustment to equilibrium. For$m$ = 30, d_y = 1.060, d_x = 1.073, and d_e = 0.688, confirming slow adjustment to equilibrium. Nevertheless, for m = 40, d_y = 1.0892, d_x = 1.156, and d_e = 0.809. Overall, the evidence for rejecting the null hypothesis in favor of market integration for beef cattle between Uruguay and Paraguay is inconclusive.

5. Discussion

The results presented above are not unexpected. Firstly, with Argentina at one extreme and Paraguay at the other, all the "individual country markets" show a general common pattern relative to the international market (Uruguay). The magnitudes of the estimated cointegration coefficients ($\hat{\beta}$) indicate that the rank ordering of the cointegrated market relationships is led by Paraguay (PY), followed by São Paulo (SP), Mato Grosso do Sul (MS), Rio Grande do Sul (RS), and Argentina (AR). However, none of the pairwise comparisons satisfied the proposed rule of thumb.

Table 5. Cointegration Analysis between Uruguay (UY) and Argentina (AR) Cattle Prices

Value of *m*	Order of Integration			Cointegration Coefficient
	UY price	**AR price**	**Error**	
m = 20	d_y = 0.774	d_x = 1.218	d_e = 0.550	β = 0.448
	CI: (0.582; 0.967)	CI: (1.045; 1.392)	CI: (0.385; 0.716)	
	p-value: 0.0008	p-value: 0.0000	p-value: 0.0037	
No. CI estimate of d_y is not included within the CI of d_x				
m = 30	d_y = 0.971	d_x = 1.089	d_e = 0.889	β = 0. 448
	CI: (0.851; 1.090)	CI: (0.992; 1.185)	CI: (0.775; 1.004)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
No. The error is non stationary (d_e> 0.75)				
m = 40	d_y = 1.033	d_x = 1.045	d_e = 0.949	β = 0. 448
	CI: (0.968; 1.098)	CI: (0.975; 1.114)	CI: (0.855; 1.043)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
No. The error is non stationary (d_e> 0.75)				

Note: 501 observations

The most dubious case was found for UY-PY. In spite of the high degree of cointegration exhibited by the correlation coefficient ($\hat{\beta}$= 0.967), the results were inconclusive in two cases (d_e< 0.75) and negative in the third. In the absence of a shock to the PY cattle market, such as the FMD outbreak of 2011-12, it is likely that the results would have tended toward rejection of the null hypothesis. There are similarities between the Uruguayan and Paraguayan markets that support the alternative hypothesis of market integration. The beef cattle sectors in the two countries have minimal public intervention and cattle prices are determined through unhindered market processes reflecting international beef prices. In recent years, both nations' beef packing industries have become increasingly concentrated, resulting in some degree of oligopsony power on the demand side of the cattle market. As noted by Goodwin & Schroeder (1991), the increasing concentration may lead to increasing cointegration. However, the world market price-taker positions of both Uruguay and Paraguay likely limits the degree of market power, which can be achieved by beef manufacturers in the two countries. Because Uruguay and Paraguay do not have common borders, a local shock (*e.g.*, changes in sanitary status, severe climate events) affecting one of them is not likely to affect the other.

Table 6. Cointegration Analysis between Uruguay (UY) and Paraguay (PY) Cattle Prices

Value of *m*	**Order of Integration**			**Cointegration Coefficient**
	UY price	**PY price**	**Error**	
m = 20	d_y = 1.044	d_x = 1.146	d_e = 0.695	β = 0.967
	CI: (0.789; 1.299)	CI: (0.980; 1.313)	CI: (0.483; 0.907)	
	p-value: 0.0007	p-value: 0.0000	p-value: 0.0042	
Inconclusive. Slow adjustment to equilibrium aftershocks (0.5 <d_e< 0.75).				
m = 30	d_y = 1.060	d_x = 1.073	d_e = 0.688	β = 0.967
	CI: (0.893; 1.227)	CI: (0.956; 1.191)	CI: (0.545; 0.832)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
Inconclusive. Slow adjustment to equilibrium aftershocks (0.5 <d_e< 0.75).				
m = 40	d_y = 1.089	d_x = 1.156	d_e = 0.809	β = 0.967
	CI: (0.942; 1.236)	CI: (1.033; 1.278)	CI: (0.691; 0.927)	
	p-value: 0.0000	p-value: 0.0000	p-value: 0.0000	
No. The error is non stationary (d_e> 0.75)				

Note: 422 observations

Alternatively, Argentina's beef cattle industry was greatly affected by macroeconomic and sector-specific policies followed by its government in recent years. These policies have had many unintended consequences. Many Argentine cattle producers have shifted from grazing cattle to growing cereal grains and soybeans because these commodities were less affected by public policies that reduce their participation in international markets. Due to national policies, Argentine beef producers lost a large portion of their global market share to beef producers in Uruguay, Brazil, and Paraguay. Furthermore, even as Argentina's beef packing sector has concentrated (similar to the other Mercosur countries), the effects of discriminatory public policies on the cattle industry have prevailed and impeded regional market integration. Under this political and economic framework, it is no surprise that market integration does not hold from the beginning ($\hat{\beta}$= 0.45) between AR and UY, as demonstrated by the empirical results.

Cointegration results between Uruguay and the three Brazilian markets lie between the UR-AR and UR-PY cases, with São Paulo ($\hat{\beta}$= 0.66), Mato Grosso do Sul ($\hat{\beta}$= 0.64), Rio Grande do Sul ($\hat{\beta}$= 0.61). The magnitude of the cointegration coefficient was relatively high, but the results showed some evidence of market integration only for SP in one case, and were inconclusive for RS and MS only in one case. In all the other cases, the results (d_e> 0.75) ruled out market integration.

Goodwin & Schroeder (1991) explained that although differences in market volume may influence trade activity across market areas, the influence on cointegration is not clear. In spite of the differences in total volume of cattle marketing between Uruguay and Brazil, differences in the proportion of total beef production traded in the domestic market relative to the export market manifested in different price formation processes between Uruguay and Brazil. Despite short run variability caused by local agro-ecological conditions, cattle prices in Uruguay fully reflect shocks occurred in the international markets. However, although Brazil is a key player in the global beef markets, the size of its domestic market buffers the effects of international market shocks.

6. Conclusions

Since its inception, Mercosur has sought to create a unified trade bloc in products for which the member nations have a comparative advantage. This study is the first to assess *actual* Mercosur beef cattle market integration.

Although the 'separate markets' hypothesis could not be rejected here, empirical results illustrate that market integration among Mercosur cattle markets is plausible at some general level, with the only exception of the Argentine market. In theory, the very existence of an economic bloc should be enough to suspect the existence of integrated markets. However, real-world circumstances are somewhat more complex. The Mercosur bloc was created in 1991 by the Asunción Treaty, and signed by the four original partners. Even after a quarter of a century, the Mercosur bloc often does not behave as a free-trade region within its borders. While broad-scale discussion of Mercosur objectives and achievements are beyond the scope of this paper, assessment of cattle market integration is relevant to understand economic conditions and prospects within that important agricultural sector and related industries throughout the region. Mercosur has likely created an advantageous framework for regional cattle market integration, as well as for improving the global market position of the bloc's comprehensive beef cattle industry. However, given regional and global trends in beef industry concentration, we are left to wonder how much of future cattle market integration will be due to industry forces (including scale economies) rather than trade bloc policy. For example, it remains uncertain the degree to which relevant Mercosur initiatives have driven or been driven by beef cattle market trends and conditions. Beyond that, the vulnerability of live cattle production to climatic, disease, and economic shocks will continue to challenge beef industry integration at the Mercosur region. Lanfranco, Castaño, Zorrilla de San Martín, & Porto (2010) suggested that, in the case of grass-fed systems, differences in Mercosur agro-ecological and climatic conditions may be the foundation of permanent spatial price differentials.

The results presented herein indicate that basic trends and changes in prices are similar in all the Mercosur cattle markets analyzed, except for Argentina. Both, beef and cattle markets in Uruguay have been subject to the signals of supply and demand of the international beef market at least in the last 20 years. In this article, the degree of integration of the former with other Mercosur markets is a measure of the degree of integration of the latter with the international market.

In Argentina, the pattern of the series is completely different from Uruguay, as the series is not balanced. They have a different order of integration, leading to the conclusion that these are two completely separate markets. The result is not surprising. Argentine government implemented a series of economic policies over the period under examination, which isolated the local economy from the world economy, and seriously affected the local cattle/beef sector.

The most significant policies were: restrictions to foreign exchange operations, limitation of beef exports, and setting beef export taxes to 15%.

The result of these policies is well-evidenced by a significant decline in cattle numbers. Argentina's beef cattle herd shrank from a high of 55.7 million head in 2007 to just 48 million in 2010(USDA-FAS, 2016). In addition, it lost presence as a relevant beef exporter. During the period 2004-2009, Argentina ranked between 3rd to 5thin the list of top beef exporters, falling to the 10th to the 12th position between 2011 and 2015 (USDA-FAS, 2016).

Large external shocks, such as dramatic movements in commodity prices in 2008, affected markets worldwide; Mercosur markets were not an exception. Global changes in supply and demand, in consumer preferences, in the industrial organization of the food industry, and more specifically meat industry characteristics from farm to the table have affected cattle markets worldwide.

However, t some dynamic forces can separately affect a particular market at a certain point in time, creating a temporary shock with variable duration and persistence. Some of these factors are market-specific, some related to particular agro-climatic and ecological conditions under which cattle production systems have developed, and some linked to the economic conditions prevailing within each country or region.

Differences in cattle production systems, including breeds and biotypes, feeding and fattening schemes, geography, tradition, and cultural aspects that determine the way cattlemen present and market the animals, may introduce noise into price discovery and market prices (Lanfranco, Ois, & Bedat, 2006). Some of these factors may result in product differentiation, which compromises the assumption of homogeneity underlying the concept of the law of one price. Variable market dynamics are most likely to manifest in regular or predictable gaps between different price series rather than in price shocks. These gaps will not preclude market integration if they are permanent and more or less proportional over time.

Other market distortions, such as those caused by climate and environmental factors or nutritional and sanitary aspects, can cause price shocks of different magnitude and persistence, especially when they involve the occurrence of extreme weather conditions or events (*e.g.*, disease outbreaks, droughts, floods). These results showed that in most cases, even when the general behavior of the price series under analysis supported the conclusion that markets are integrated, price responses to specific shocks were dissimilar and, most importantly, the reversion of prices to equilibrium was always very slow.

As noted above, market integration is a moving target. Future applications of the methodology outlined here will aid in assessing Mercosur region progress toward integration of cattle and other markets. These results for Mercosur cattle prices are consistent with cointegration studies that have used techniques based on full differencing price series (Fossati, Lorenzo, & Rodríguez, 2007). They are also consistent with the results obtained by Mohanty *et al.*, (1998), who examined price integration in the MERCOSUR countries of Brazil and Argentina for wheat and corn, using an augmented Dickey-Fuller (ADF) and the GPH tests.

The methods applied here also provide additional information that the earlier studies were unable to generate. The analytical procedures applied to Mercosur cattle prices have the potential to provide new information about market integration processes and progress within and between other regions and nations. Assessing degrees of market integration improves accountability for entities tasked with removing barriers to trade and harmonizing trade between regions, nations, and blocs. These evaluations also provide insight into market integration processes which may be occurring in spite of governmental or treaty-driven efforts rather than because of them.

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