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Beef Cattle in the MERCOSUR bloc: Integrated or Separate Markets?

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Abstract

This aim of this paper was to provide empirical evidence of market integration in the beef cattle industry of the MERCOSUR economic bloc. This region possesses the largest commercial bovine herd in the world and explains one quarter of the international trade for beef. The study included six relevant cattle markets located in the four original members of the bloc. The analysis was conducted using a fractional cointegration approach proposed Marinucci and Robinson (2001). The fractional integration parameters and the error term of the cointegration equations were computed using a semi-parametric approach developed by Geweke and Porter-Hudak (1983). The null hypothesis of 'separate markets' could not be rejected even when the general behavior of the series under analysis encouraged the idea of some market integration. Despite the common trends, their response to specific shocks was dissimilar and, most importantly, the reversion of the disturbed series to equilibrium was always slow.

Keywords: cattle prices, law of one price, market integration, fractional cointegration.
JEL codes: C5, F15, Q11

1. Background

1.1. Cattle production in the MERCOSUR bloc: Integrated or separate markets?

With more than 290 million bovine heads, the MERCOSUR (Brazil, Argentina, Paraguay, and Uruguay) region possesses about 20% of the total world's herd and is one of the most important beef producing regions on the planet. Taken individually, Brazil holds the largest commercial herd in the entire world (217.4 million heads). In turn, Argentina and Uruguay are well known globally as producers of high quality beef, while Paraguay has exhibited an important growth in beef production in recent years. One of the distinctive characteristics of cattle production in this region is the absolute predominance of open-sky grass-fed production systems.

Together, the four members of the economic bloc produced 12.7 million metric tons (MMT) of bovine meat (carcass weight) in 2012 (**Table 1**), accounting by 20% of total world’s production. The MERCOSUR inhabitants are heavy beef eaters. Argentinians and Uruguayans dispute the first place in the ranking of the world's of per capita consumption of this meat (Romero, 2013). Nevertheless, the region is also a major supplier to the rest of the world. All of them, Brazil, Uruguay, Argentina, and Paraguay rank in the list of top beef exporters (USDA, 2014). In 2013, the bloc exported 1.8 MMT (shipped weight) to 120 different markets, for a total FOB value of US$ 8.8 billion.

<TABLE 1>

There are some similitude, but also important differences among the diverse cattle production systems found in this region, which integrates dissimilar geographic and agro ecological conditions. In tropical and subtropical areas, cattle production is based on zebu herds (*Bos taurus indicus*). Over 80% of the Brazilian herd belongs to this type. Nelore is the predominant breed representing 90% of zebu cattle in this country (Duran, 2014). Zebu breeds and crossbreeds have great resistance to heat. They exhibit great capacity of adaptation and had proven to be very efficient in producing beef in such difficult conditions. Zebu crossbreeds mainly Brangus (Brahman×Angus) and Braford (Brahman×Hereford) are common in the subtropical areas of Brazil and Paraguay, and, in less extent, in the north of Argentina and Uruguay.

On the other hand, the "taurine" type of cattle (*Bos taurus taurus*) prevails in the more tempered areas of the South American continent region. These areas, broadly delimited by the Río de la Plata River basin, include grassland areas in the southern states of Brazil (Santa Catarina, Paraná, and especially Rio Grande do Sul), in both the Pampean and the Mesopotamic provinces of Argentina (Buenos Aires, Entre Ríos, Santa Fe, and part of Corrientes), in Uruguay, and in the south of Paraguay. The predominant beef cattle breeds, either pure or crossbreeds, are the British breeds (Angus, Hereford, Shorton) although the continental breeds (Charolais, Limousin) have also their own space.

This area is of especial interest as 'taurine' cattle and especially British breeds are very appreciated in the cattle markets because of the quality of the beef they can produce. In addition, under adequate conditions of climate and nutrition, this type of cattle is very productive and very efficient in converting feed to muscle and, thereafter, to meat. Other things equal, fattened cattle from these breeds receive better prices than other cattle types (zebu and even continental breeds) at the time of slaughter.

An important question arises at this point. Can this vast cattle producing region, probably the most relevant from the point of view of commercial livestock farming, be considered as one big, spatially separated but still single integrated market or, by the contrary, it comprises several separate markets, each one with its own peculiarities and completely independent one to another? If so, to which extent are they integrated?

The present study is an attempt to offer some answer to this question, at least partially. In order to measure the degree of integration of the diverse cattle markets operating within the MERCOSUR region, this research relies upon the theoretical concept behind the "law of one price". As pointed out by Ravallion (1986), measurement of market integration can be viewed as basic data for understanding how specific markets work. The extent to which commodity markets are integrated also has important implications for government regulation and general economic policy (Fossati, Lorenzo, and Rodríguez, 2007).

Some issues can be argued in support of one or the other hypothesis. As pointed out before, Brazil, Argentina, Paraguay, and Uruguay are all members of the MERCOSUR bloc, whose name means precisely "Common Market of the South". In theory, this might be a powerful argument in favor of the hypothesis of one single market.

According to Article 1 of the so-called Treaty of Asunción, which gave birth to MERCOSUR, the main objectives[1] of the bloc are: a) The free movement of goods, services and factors of production between the countries, through, inter alia, the elimination of customs duties and non-tariff restrictions on the movement of goods and any other equivalent measure; b) the establishment of a common external tariff and the adoption of a common commercial policy in relation to third States or groups of States and the coordination of positions in economic forums regional and international trade; c) the coordination of macroeconomic and sector policies among States Members: foreign trade, agricultural, industrial, fiscal, monetary, exchange rates and capital of services, customs, transport, and communications and other agreed upon, in order to ensure adequate conditions of competition between the States Members; 4. the commitment of States Members to harmonize their legislation in the relevant areas in order to strengthen the integration process.

Another point in favor of the one-market hypothesis is the strong globalization faced by the beef industry in its manufacturing sector (slaughter, processing, and packing). The most notorious case in that sense was the international expansion of Brazilian firms. Bittencourt, Carracelas, and Reig (2011) stated that since 2005, the two largest economic groups of this

[1] http://www.mercosur.int/t_generic.jsp?contentid=3862&site=1&channel=secretaria&seccion=2. Last Access: Oct 2014.

origin, JBS and Marfrig, acquired beef plants in South America, United States, Europe, and Australia. Only in the MERCOSUR region, the JBS group purchased more than ten plants in Argentina, while the Marfrig group acquired four plants in Argentina and five in Uruguay. This means that the most relevant players on the demand side of the beef cattle business are the same in the MERCOSUR region, no matter the country.

On the other hand, there are factors conspiring against the idea of a single market. First, in practice, the existence of the MERCOSUR as an advantage of an integrated cattle market depends on the actual level of development of the economic bloc. In fact, many of the MERCOSUR objectives are still far from being accomplished; in particular, those referred to coordination of macroeconomic and sector policies, and the movement of goods, services and production factors between the members of the bloc.

Second, as discussed before the agro-ecological and geographic conditions put restrictions in the way production develops in a certain area, including the type of cattle used. The characteristics of the product raised in such conditions may be quite different with consequences in terms of market value. Cattle markets can be viewed as differentiated markets, even within the same geographic location due to differences in the endowment of animal traits (Buccola and Jessee, 1979; Buccola, 1980).

Lanfranco *et al*. (2010) suggested that, in the case of grass-fed systems, differences in agro-ecological and climatic conditions derive in permanent price differentials between geographic zones. The type of cattle can, by itself, be another source of a price differential. Lanfranco, Ois, and Bedat (2006) found differences in the price received by lots of feeder cattle because of their predominant breed or crossbreed, as well as other phenotypic traits, *ceteris paribus*.

1.2. Market integration and the Law of One Price

The "Law of One Price" (LOP) implies that, for a given commodity, a representative price will prevail across geographically separated markets located in one or more countries, after it is adjusted by exchange rates and allowance for transportation costs. The LOP defines the extent of the market and measures the degree of integration among markets. Deviations from LOP can be explained by factors, such as the short run volatility of exchange rates and other "overshooting effects" (Ardeni, 1989).

The basic idea behind the spatial market integration is discussed by Takayama and Judge (1971). This theory develops a model under the assumption of free flow of information and goods, prices of a homogeneous good in two spatially separated markets should only differ by the transaction costs. If the price in one market is larger than the price in another market plus the transaction cost that would be involved if one had to move the product from the market with the low price to the market with the high price, unexploited pure profits would exist.

Rational traders would therefore enter the market and capitalize on these arbitrage opportunities, increasing demand in the market where prices are low and increasing supply in the location where prices are high. These two forces will, *ceteris paribus*, drive up the price in the market that had initially a low price and reduce the price in the market that had initially a high price. At the end, prices adjust up to the point where trade only derives in normal profits. That is, the price difference becomes equal to the transaction cost.

Therefore, price changes in the export region induce price changes in the import region, in the same direction and magnitude. If this is the case, the two markets are completely integrated as a single market. The extent and the speed to which shocks are passed through, and the strength of the interdependence among prices are indicators of the degree of integration and global efficiency of markets' performance.

Among the different methods used to measure market integration, cointegration analysis is a widespread approach. Part of its appeal for testing LOP comes from the fact that most of price series are non stationary, and the theory of cointegration allows for the testing long run relationships between or among economic variables in the presence of non stationary.

This paper examines the LOP using the fractional cointegration approach proposed by Marinucci and Robinson (2001). The fractional integration parameters and the error term of the cointegration equations are estimated using a semi-parametric approach developed by Geweke and Porter-Hudak (1983). Fractional 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.

Thus, 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. In addition, the variables can be analyzed in levels, without any transformation. This reduces the loss of information that arises during each transformation when using standard methods.

The LOP assumes that the prices of goods are geographically arbitraged and, adjusted for tariffs and transport costs, there are equalized among locations. This formulation ignores differences in product qualities (homogeneity) and in transportation, storage, and marketing costs, as well as other domestic non tradable inputs.

Let P_t , P_t^*, and e_t to denote respectively the price of a certain good in the "home country" and expressed in the local currency, the foreign price expressed in its own currency, and the exchange rate in terms of home currency price per unit of foreign currency, all at moment t. Arbitrage can be reached by $P_t = e_t.P_t^*$. An important implication of complete spatial arbitrage is the idea that relative national price levels in a common currency are independent of the exchange rate, since exchange rate movements merely reflect, passively, divergent national trends. This is an application of the homogeneity postulate which holds when money is fully neutral (Dornbush, 1987).

To allow for deviations from this assumption, as well as for effects not included in the model, a disturbance term denoted by u_t can be added to the equation in order to estimate the following regression:

$$\frac{P_{it}}{e_{it}} = \beta . P_{it}^{*} + u_{it} , \qquad i = 1,\ldots, p.$$

The coefficient β is the parameter of fractional cointegration developed for Marinucci and Robinson (2001). Values close to 1 imply that a variation in prices is fully transmitted to the domestic prices, whereas a value of 0 implies no transmission at all. On the other hand, the value of the fractional integration parameter of the error refers to the velocity of adjustment to the equilibrium.

2. Data and Methods

2.1. Empirical Approach

2.1.1. Fractional Cointegration

Let $\mathbf{x}_t$, be a vector of economic variables with t as a time subscript. The variables are in equilibrium when the linear relationship type $\alpha'\mathbf{x}_t = 0$. Most of the time, the variables in $\mathbf{x}_t$ will not be in equilibrium, and the departure from equilibrium can be defined as $\mathbf{z}_t = \alpha'\mathbf{x}_t$. The components of vector $\mathbf{x}_t$ are cointegrated of order d, b, which is expressed as $\mathbf{x}_t \sim \mathrm{CI}(d,b)$, if: (a) all the components of $\mathbf{x}_t$ are integrated of d. (b) there is a non-zero vector α such that $\mathbf{z}_t = \alpha'\mathbf{x}_t \sim \mathrm{I}(d\text{-}b)$, for $b > 0$. Vector α is then called as the cointegration vector (Engle and Granger, 1987).

Given the following multivariate equation $y_t = \mathbf{x}_t'\beta + e_t$, Given the vector $\mathbf{z}_t = (\mathbf{x}_t', y_t)$, where y_t is a scalar and $\mathbf{x}_t = (x_{(1)},\ldots, x_{(p\text{-}1)t})'$, it is said that $\mathbf{z}_t$ is cointegrated (of orders $d_1,\ldots,d_{p\text{-}1}$, d_y;

$d(\beta)$) if x_{it} is I(d_i), with $i = 1,\ldots, p$-1, and y_t is I(d_y), and there exist a vector $\beta_{(p-1)\times(1)}$, such that $e_t = y_t - \beta'\mathbf{x}_t$ is I[$d(\beta)$], for $d(\beta) < d_y$ (Robinson and Marinucci, 2001).

Fractional cointegration for a $p\times1$ dimension vector $\mathbf{z}_t$, whose i-th element is $\mathbf{z}_{it} \equiv$ I(d_t), $d_i >$ 0, $i = 1,\ldots, p$, where $\mathbf{z}_t \equiv$ FCI($d1,\ldots, d_p, d_e$) if there exist a non-zero $p\times1$ dimension vector α, such that $e_t = \alpha'\mathbf{z}_t \equiv$ I(d_e), for which $0 \leq d_e <$ min1 $\leq i \leq pd_i$ (Robinson and Marinucci, 1998). This property is possible and makes sense if and only if $d_i = d_j$, for some $i \neq j$. A necessary condition for α to be a cointegration vector is that the i-th component be equal to zero if $d_i > d_j$ for all $i \neq j$. In the case when $d_1 = d_2 =\ldots= d_p = d$ it is common to write $\mathbf{z}_t \equiv$ CI(d,b), where $b = d - d_e$ measures the intensity of the cointegration relationship. The cointegration vector defined by Engle and Granger (1987) is then a particular case and it is denoted $\mathbf{z}_t \equiv$ FCI(1,…1, 0) or analogously CI(1,1).

The approach of Engle and Granger (1987) has been followed in the economic literature due to different reasons. First, unit roots can be seen as a consequence of economic theory, for example, the hypothesis of market efficiency, and the random walk hypothesis of consumption. Second, the usual standard tests tend to fail in favor of the rejection of the null hypothesis of unit root in many time series. Third, the computational implications of the unit root hypothesis, which allows the differentiation of the series to remove the non-stationarity, are attractive. Four, the asymptotic theory for the statistics based on sequences I(0) have been better developed than those for stationary processes I(d), with $d > 0$. Finally, inference rules related to processes of fractional stationarity and non-stationarity I(d) have not yet been well developed.

2.1.2. Long Memory Processes

From the empirical point of view, the concept of long-term memory is usually related with the persistence showed by the autocorrelations sample of certain stationary time series, which converge to zero at a very slow pace. This behavior is not compatible with the autocorrelation functions of the models of autoregressive moving average or ARMA, which impose an exponential decrease of the autocorrelations, nor with the degree of persistence in the autoregressive integrated moving average models, ARIMA (Castaño, Gómez, and Gallón, 2008).

When building a model, the usual practice of differentiating the series to achieve stationarity may have negative consequences (Granger and Joyeux, 1980). The problem arises when many apparently non-stationary economic series with spectrum not bounded in the origin are differentiated to achieve a finite variance. The differentiated series can be converted into a series with null spectrum in the zero, indicating that the low frequency component, which is very important in the long term forecast, has been removed from the original series. 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, some authors (Granger and Joyeux, 1980; Granger, 1980; Hosking, 1981) introduced the so-called ARFIMA models (autoregressive fractionally integrated moving average) for modeling economic series. The ARFIMA models cover the gap between the extreme cases of unitary root models, and stationary models that impose exponential decrease of the autocorrelations and therefore a spectrum bounded in the zero frequency. They are models of autoregressive moving average where the differentiation is fractional. The differencing parameter d is not an integer but a real number. These models cover the "intermediate case" that exists between the unitary root ARIMA processes and the ARMA processes (Pérez, 2001). The ARFIMA processes produce long memory if the parameter of differentiation is in the range $0 < d < 1/2$, in which case the process is stationary and invertible.

On the other hand, while for integrated processes ($d = 1$) the effect of a shock persists indefinitely, in a fractionally integrated process with $0 < d < 1/2$, the effect of a shock just ends up disappearing 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 cancelled in the origin and it is dominated by the high frequencies, while the autocorrelations are all negative and absolutely summable. 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 can not do because the process will be non stationary.

Formally, an ARFIMA process $\{y_t\}$ is defined by

$$\Phi(B)y_t = \theta(B)(1-B)^{-d} + \varepsilon_t\,, \tag{1}$$

where $\Phi(B) = 1+\phi_1 B+\ldots+ \phi_p B^p$ is the autoregressive operator and $\theta(B) = 1+\phi_1 B+\ldots+ \phi_q B^q$ el is the moving average operator. The lag operator B is used instead of L, since the differences are fractional; $\Phi(B)$ and $\theta(B)$ do not have common roots and $(1 - B)^{-d}$ is the fractional differencing operator defined in the following way:

$$(1-B)^d = \sum_{j=0}^{\infty} \frac{\Gamma(j-d)}{\Gamma(-d)\Gamma(j+1)} B^j\,, \qquad \Gamma(a) = \int_0^{\infty} x^{a-1} e^{-x} dx\,. \tag{2}$$

ARFIMA processes allow to simultaneously describing the dynamic properties in the long term through the parameter d, and the short-term correlation through the parameters p and q of the ARMA part of the model.

2.1.3. Time and Spectral Domains

Any stationary and invertible time series can be represented by the Wold Theorem, as

$$y_t = \sum_{j=0}^{\infty} \psi_j \varepsilon_{t-j}\,. \tag{3}$$

It focuses on the implications of the covariances at different moments in time, analyzing the properties of $\{y_{t_t}\}_{t=-\infty}^{t=+\infty}$ in the domain of time (Hamilton, 1994).

A different approach consists in expressing the variable y_t as a weighted sum of trigonometric functions. This approach is called spectral analysis of time series or frequency domain, which is an extension of the Fourier method (Pollock, 1999). According to the theorem of the spectral representation (TSR), the series can be expressed as follows:

$$y_t = \sum_t \left[\alpha_j \cos(\lambda_j t) + \beta_j \mathrm{sen}(\lambda_j t)\right] + \varepsilon_t \ , \tag{4}$$

where ε_t is white noise, independent and identically distributed (IID); α_j and β_j are the Fourier coefficients. A formal derivation of the TSR for finite samples can be stated by letting y_{t_t} to be a stationary process with absolutely summable autocovariances so that $\sum_{j=0}^{\infty} |\gamma_j| < \infty$,. Given a sample of size T, with realizations $\{y_1, y_2,\ldots, y_T\}$, the $T-1$ sample autocovariances can be calculated as:

$$\gamma_j = T^{-1} \sum_{t=j+1}^{T} (y_t - \bar{y})(y_{t-j} - \bar{y}) \qquad \text{for } j = 0, 1, 2, 3, \ldots, T\text{-}1, \text{ and}$$

$$\gamma_j = \gamma_{-j} \qquad \text{for } j = \text{-}1, \text{-}2, \text{-}3, \ldots, \text{-}T\text{+}1, \tag{5}$$

being the sample mean,

$$\bar{y} = T^{-1} \sum_{T=1}^{T} y_t \ . \tag{6}$$

For any λ, the sample periodogram can be expressed as:

$$\hat{I}_y(\lambda) = \frac{1}{2\pi}\left[\hat{\gamma}_0 + 2\sum_{j=1}^{t=1} \hat{\gamma}_j \cos(\lambda_j)\right]. \tag{7}$$

The sample variance of y_{t_t} is:

$$\int_{-\pi}^{\pi} \hat{I}_y(\lambda) d\lambda = \hat{\gamma}_0 \, . \tag{8}$$

The sample periodogram is symmetric for any $\lambda = 0$. In a equivalent way, it can be written as:

$$\hat{\gamma}_0 = 2\int_0^{\pi} I_y(\lambda)d\lambda . \tag{9}$$

Analogously for a sample process, the TSR can be expressed as:

$$y_t = \hat{\mu} + \sum_{j=1}^{M} \left\{ \hat{\alpha}_j \cos\left[\lambda_j(t-1)\right] + \hat{\delta}_j sen\left[\lambda_j(t-1)\right] \right\}. \tag{10}$$

The sample variance of y_t and the portion of this variable can be attributed to cycles with frequency, λ_j that can be inferred from the sample periodogram $\hat{I}_y(\lambda)$. In this way, it is possible to estimate a behavioral model for a time series in the domain of Fourier frequencies. Equation (10) can be viewed as a standard regression where the independent variables are orthogonal. The Fourier coefficients can be estimated by OLS using the following expression:

$$y_t = \beta' \mathbf{x}_t + \mu_t. \tag{11}$$

2.1.4. Estimation of the long-memory parameter

In this research, the chosen method for estimating the long-memory parameter d, is the semi-parametric method proposed by Geweke and Porter-Hudak (1983), known as GPH. The problem to be solved is the estimation of parameter d in the following expression:

$$(1 - B)^d y_t = \mu_t . \tag{12}$$

Geweke and Porter-Hudak (1983) demonstrated that if y_t is an integrated series with memory parameter d, such that $-1/2 < d < 1/2$, this relationship actually holds if and only if there exist a fractional Gaussian white noise whose parameter is $H = d + ½$, and u is a stationary linear process with spectral density function $f_\mu(\lambda)$ that is finite and continuous in the interval $[-\pi,\pi]$. The spectral density function of y_t is defined by:

$$f(\lambda) = \left(\frac{\sigma^2}{2\pi}\right)\left\{4\text{sen}^2(\lambda)\right\}^{-d} f_\mu(\lambda). \tag{13}$$

Taking logs and assuming a sample of size T for y_t, arises the following expression:

$$\log I(\lambda_j) = a - d\log\left\{4\text{sen}^2\left(\frac{\lambda_j}{2}\right)\right\} + e_j, \qquad j = 1, 2, \ldots, m. \tag{14}$$

The term $I(\lambda_j)$ is the periodogram computed in Fourier frequencies $(\lambda_j) = 2\pi_j / T$:

$$I(\lambda_j) = \frac{1}{2\pi . N}\left|\sum_{t=1}^{N} x_t e^{it\lambda_j}\right|^2, \qquad j = 1, 2, \ldots, m.$$

The parameters of the regression, including d, can be estimated by OLS. Robinson (1995) demonstrated the consistency and the asymptotic normality for the range $-½ < d < ½$.

2.1.5. Estimation of the Fractional Cointegration Vector

The first step for computing the fractional cointegration vector is the estimation of β in the representation $y_t = \beta'\mathbf{x}_t + e_t$ for the observed vector $\mathbf{z}_t = (\mathbf{x}_t', y_t)$. The e_t process is $I(d_e)$ with $d_e < d_y$, assuming that β is identified. It is also assumed that $\mathbf{z}_t$ is observed for $t = 1, \ldots, n$. Thus, the following discrete Fourier transformation is defined for vector $\mathbf{y}_t$, $t = 1, \ldots, n$.

$$w_y(\lambda) = \frac{1}{(2\pi n)}\sum_{t=1}^{n} y_t e^{it\lambda} . \tag{15}$$

The cross-periodogram is defined for the vector or scalar sequence x_t, $t = 1, \ldots, n$:

$$I_{yx}(\lambda) = w_y(\lambda) w_x(-\lambda). \tag{16}$$

Denoting $\lambda_j = 2\pi\, j/n$ for the integer j of the Fourier frequencies and defining the mean cross- periodogram as:

$$\hat{F}_{yx}(m) = 2\,\text{Re}\left\{\frac{2\pi}{n}\sum_{j=1}^{m} I_{yx}(\lambda_j)\right\} - \frac{2\pi}{n} I_{yx}(\pi)\ell\left(m = \frac{n}{2}\right), \tag{17}$$

where $\ell(\cdot)$ is the index function and the integer m is the bandwidth of the periodogram, and satisfies $1 \le m \le n/2$. In turn, Re stands for the real part of the expression between brackets.

The second term of (17) only makes a contribution when m attains its maximum value in $n/2$. The function $\hat{F}_{yx}(m)$ represents the contribution of frequencies $[1,\lambda_m]$ to the sample variance.

Assuming that the inverse does exist, the parameter β is estimated by OLS in the frequency domain,

$$\hat{\beta} = \hat{F}_{xx}(m)^{-1}\hat{F}_{xy}(m). \tag{18}$$

Robinson (1995) proposed (19) when (p = 2), with stationary series x_t and y_t. If (17) is taken as a special case, the estimation de β by OLS with an intercept will be,

$$\hat{\beta}_{n/2} = \left[\sum_{t=1}^{n}(x_t - \bar{x})(x_t - \bar{x})'\right]^{-1}\sum_{t=1}^{n}(y_t - \bar{y})(x_t - \bar{x})'. \tag{19}$$

When $m < n/2$, there exist two cases of interest in the asymptotic context, where $n \to \infty$,

$$m \sim C.n \qquad \text{with } 0 \le C \le ½, \text{ and} \tag{20}$$

$$\frac{1}{m} + \frac{m}{n} \to 0. \tag{21}$$

Considering the total Fourier frequencies of $m = n/2$, the result of the estimation of the cointegration vectors is the OLS procedure. The sample distribution of these estimators does not lie within any of the known tabulated distributions. However, Marinucci and Robinson (2001) derived a distribution in the limit of $\hat{\beta}_m$ using Monte Carlo simulation considering four different situations, for the rate of convergence of the estimator.

2.2. Data

The cointegration analysis was performed by considering Uruguay as the “home country” and Brazil, Argentina, and Paraguay as the foreign countries. Thus, all the comparisons are made against this market. A number of reasons were considered for taking this approach.

First, the sanitary status of Uruguay is singular. While the four MERCOSUR 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 (MGAP, 2014; OIE, 2014). The only other country that shares the same status in the world is South Korea. The other partners of the bloc, Brazil, Argentina, and Paraguay have each of them one or more zones declared FMD with vaccination. However, none of them enjoys this status as a country.

Second, setting aside Paraguay, who is a relatively new player in the global beef market, Uruguay exports the largest proportion of the beef produced in the country, among the other three (Table 1). It has an important tradition as a beef exporter and its performance in the international market. In Uruguay, both beef and cattle are quoted and traded in US dollars. Moreover, beef exporters as well as cattle producers receive payments in that currency when selling their product. Setting Uruguay (UY) as the "home country" for the analysis gives the advantage of using the US dollar directly as the local currency and avoiding the arbitrage of two currencies using a third one. The only relevant exchange rate is that of the US dollar (local) with Brazil, Argentina, and Paraguay, taken individually.

In turn, given the heterogeneity of productive regions in Brazil, this country was not considered, a priory, as a single market. The states of Sao Paulo, Mato Grosso do Sul, and Rio Grande do Sul were included in this study as separate markets. This selection was made on merit to their relevance as cattle producers and their proximity to the other markets under analysis. Together, these three states concentrate 22% of the national bovine herd and 40% of the beef exports of the country.

The price series were obtained from local sources in each country. Uruguay (UY) prices were provided by the *Instituto Nacional de Carnes* (INAC). The 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. In turn, Argentina (AR) prices correspond to Liniers cattle market (*Mercado de Liniers S.A.*); the price series for Paraguay (PY) was obtained from a private database provider.

As pointed out by Rostán (2009), a precise computation of periodograms and cross-periodograms, in the way it was done in this research, is only possible when using large series of data, with at least 300 observations. The analysis was conducted using average weekly prices for fat steers ready for slaughter, quoted in US dollars per kilogram, carcass weight, spanning a period of 501 weeks (observations), starting in the third week of Apr 2003 and finishing in the last week of Nov 2012. In the case of Paraguay, the available dataset comprised a shorter period, going from first week of Nov 2004 to last week of Nov 2012. For that reason the analysis of the pair UY-PY was done with only 421 observations.

3. Results and Discussion

3.1. Rule of Thumb

Robinson and Marinucci (1998) warned that no inference rules had been well developed for fractional stationary and non-stationary processes. The lack of a formal standardized test for market integration through a fractional cointegration analysis compels for following an *ad hoc* procedure and setting an arbitrary rule of thumb. The analysis is set in pairs, the home market against a foreign market, so that the same procedure is followed for each pair of price series (UY-SP, UY-RS, UY-MS, UY-AR, and UY-PY). The null hypothesis (Ho) stands for separate markets, that it, home and foreign markets are not cointegrated. The alternative hypothesis (Ha) proposes that both markets are cointegrated.

The first step consists in the visual inspection of the series through their graphical representation. This provides a good idea about their general behavior and how well they evolve each one respect to the others. By observing the evolution of the series over time it is possible to associate abrupt movements that may have occurred due to local and external shocks in the sector or in the whole economy, facilitating the understanding of the results.

In all cases, the statistical analysis was carried out for three different values of m (20, 30 y 40), as it was defined in (17), to improve the level of security for interpreting the empirical results. A pair of series was considered cointegrated if the result was positive for at least two values of m. Performing the analysis, one-by-one, for each pair of series (UY against each of the others), the vector of economic variables $\mathbf{x}_t$, is univariate and, therefore a scalar (x_t), as well as the coefficientβ. The magnitude of β offers a measure of the cointegration between the series. In addition, the reported results also include the estimated orders of fractional differencing of the 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 study, each pair of price was considered as exhibiting such balance if the point estimate of d_y was included within the 90% CI of d_x and vice versa. If it 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, so that they can not be fractionally integrated.

Next, the differencing parameter of the error term is compared with the integration orders of the series under analysis. The value of d_e has to me smaller than the value of d_y and d_x. In addition, the behavior of the error term in the cointegration relationship has to be analyzed. For an unambiguous conclusion that the series are fractionally cointegrated, the error term must be stationary ($d_e < 0.5$), If the error is between 0.5 and one ($0.5 < d_e < 0.75$), it means that it reverse to its mean and doubts about the cointegration relationship arise. It is because although it goes back to its equilibrium after a shock, the velocity of this adjustment is very slow and may take a long time.

3.2. Visual Inspection of the Series

Prior to the statistical analysis, each pair of series was compared graphically for a primary evaluation. Due to the lack of space, not all the comparisons were included in this article. Instead, and with the objective of facilitating the visual inspection, two separate graphs are presented as an example. In **Graph 1**, the price series for Uruguay (UY) is simultaneously compared with São Paulo (SP) and Rio Grande do Sul (RS). In **Graph 2**, UY prices are compared with Argentina (AR) and Paraguay (PY), although in this case, considering only the last 260 weeks of the period. Showing a similar behavior than the other Brazilian cattle markets, Mato Grosso do Sul (MS) was not included in the example. Nevertheless, this omission does not preclude from visualizing the whole picture for the analysis.

The first thing to note is that, in general terms, cattle prices follow a similar pattern over time, especially when considering large periods. However, there are some important differences among them when shorter periods are taken. For example, comparing UY, SP and RS (Graph 1), it is observed that all responded in a similar way before the boom of commodity prices in 2008 and their subsequent fall. The three series grow up rapidly since the first semester of 2007 until their collapse in Oct 2008. However, the fall was deeper for UY than for SP and UY, probably because Brazil has an important domestic market for beef that acted as a buffer for cattle prices in this country. On the contrary, Uruguay is more exposed to shocks in the international markets. While Uruguay exhibits the highest per capita consumption of beef, the local market is very small due its reduced population.

<GRAPH 1>

A couple of interesting issues arise from comparing UY, AR, and PY. First, the behavior of Argentina has clearly decoupled from UY and PY, which observe a much more similar pattern between them. This also has to be with the relevance of the domestic market in Argentina with respect to exports, as well as with the important differences of macro and sector policies applied in this country. On the contrary, Uruguay and Paraguay already export more than two-third of their beef production and government intervention in these sectors is virtually inexistent, at least compared to Argentina. Thus cattle prices in UY and PY follow international prices, and therefore suffer the effects of global shocks more closely than AR.

However, and as a second point to highlight here, internal shocks occurring in a specific country can have important consequences. This is what happened after the last FMD outbreak in Paraguay, in September 2011. Cattle prices went down sharply in this country as beef exports were suspended. Moreover, in addition to the amount of time without new foci of disease that a country must wait in order to recover its previous sanitary status, the last outbreak occurred in October 2012, well at the end of the period under analysis.

<GRAPH 2>

The examples discussed in this section highlight the importance of the visual inspection of the series under study to understand the results of the empirical results.

3.3. Empirical Results

The results of the UY-SP analysis are presented 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 would be fractional cointegration. The orders of differentiation of the fractional series taken in levels were not significantly different between them, as the corresponding CI overlap; the cointegration error was less than 0.5 and lower in magnitude to the orders of the series their in levels.

<TABLE 2>

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 non stationary, suggesting no cointegration between the prices. In accordance with the *ad hoc* rule of thumb defined in this research, there is not enough evidence to reject the null hypothesis, meaning that the cointegration between cattle prices of UY and SP could not be established.

The analysis for UY-RS is reported 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. This means that, 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 and, in addition, the error is non stationary. According to the rule of thumb, the statistical evidence does not allow to suggest cointegration between cattle prices of Uruguay and Río Grande do Sul.

<TABLE 3>

A similar outcome was found for UY-MS comparison presented in **Table 4**, 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, meaning 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 markets Uruguay and Mato Grosso do Sul.

<TABLE 4>

The results for UY-AR that appear in **Table 5**, were more conclusive in terms of accepting the null hypothesis than the results of the previous analysis. In this case, $\hat{\beta}$ was 0.448 showing 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. The error integration order was greater than 0.5 but smaller than 0.75, suggesting a very slow reversion to equilibrium after disturbing shocks. However, for m = 30, d_y = 0.971; d_x = 1.089; and d_e =

0.889, revealing a non-stationary error. The same for m = 40, where d_y = 1.033; d_x = 1.045 and d_e = 0.950. As before, following the rule of thumb, no evidence is found whatsoever to suggest the integration of Uruguay and Argentina cattle markets.

<TABLE 5>

The result for UY-PY are depicted in **Table 6**. The magnitude of the estimated cointegration coefficient was the highest among all the comparisons ($\hat{\beta}$ = 0.967), showing that both series follow similar pathways, except at the end of the period due to FMD outbreak in Paraguay. As pointed out before, the UY-PY analysis was made with 422 observations. In other words, the period under analysis was shorter than for all the other cases by 79 weeks. Thus, the computed d_y for UY prices was different in this case, for all the values of m.

With m = 20, the integration orders were d_y = 1.043; d_x = 1.146 and d_e = 0.69 the result is a slow adjustment to the equilibrium after a shock. For m = 30, d_y = 1.060; d_x = 1.073; and d_e = 0.688, confirming low adjust to the equilibrium. Nevertheless, for m = 40, d_y = 1.0892; d_x = 1.156; and d_e = 0.809. Overall, the evidence in order to reject the null hypothesis in favor of market integration for beef cattle between Uruguay and Paraguay is inconclusive.

<TABLE 6>

3.4. Discussion

The results obtained so far in this study are not unexpected. In the first place, with Argentina in one of the extremes and Paraguay in the other, all the “foreign markets” show a general common pattern in comparison to the home market (Uruguay). According to the magnitudes of the estimated cointegration coefficients ($\hat{\beta}$), the order of this relationship, in decreasing order puts Paraguay (PY), São Paulo (SP), Mato Grosso do Sul (MS), Rio Grande do Sul (RS), and Argentina (AR). However, no one of the relationships satisfied the proposed rule of thumb.

The most dubious case was found with UY-PY. In spite to the high cointegration shown by the correlation coefficient ($\hat{\beta}$ = 0.967), the results were inconclusive in two of the cases (d_e < 0.75) and negative in the remaining one (d_e > 0.75). It is likely that in the absence of a shock in the PY cattle market, such as the FMD outbreak of 2011-12, the outcome could be more promising towards the rejection of the null hypothesis.

There are some similarities between these two markets that come in support to the alternative hypothesis of market integration. The beef cattle industry operates with minimum or not public intervention in both countries and prices are quoted in open markets. In addition, both sell a big proportion of their beef production on the international market so that, in general, cattle prices mostly reflect international beef prices.

In the opposite direction, it can be argued that in recent years, both countries have experienced in increase of concentration in the industrial phase of the chain, deriving in oligopsony power on the demand side of the cattle market. However, both are price takers in the international market so that the effectiveness of a potential market in the beef manufacturing process is at least limited. However, both markets are geographically separated. Uruguay and Paraguay do not have common borders. Thus, a local shock (sanitary, climate) affecting one of them is not likely to affect the other.

On the opposite side, the beef cattle industry in Argentina is strongly affected by the general and specific policies followed by its government, at least in the last decade. These policies have derived in several unintended effects. Many cattle producers have shifted from grazing cattle to grow cereal grains and soybeans. These commodities have also been targeted by public policies that, overall, have been less disruptive in terms of their participation in the international markets.

In the case of the beef sector, the net outcome was clearly negative in that sense, as Argentine beef producers gave up a good portion of their foreign market share, many of which were seized by Uruguay, Brazil, and Paraguay have seized. Under this political and economic framework, it is not a surprise that market integration does not hold from the beginning ($\hat{\beta}$ = 0.45) between AR and UY, as it was shown by the empirical results.

Between these two extreme cases, the outcome emerged when comparing Uruguay with the three Brazilian markets, São Paulo ($\hat{\beta}$ = 0.66), Mato Grosso do Sul ($\hat{\beta}$ = 0.64), Rio Grande do Sul ($\hat{\beta}$ = 0.61), lye in the middle. The magnitude of the cointegration coefficient was relatively important, but the results showed some evidence of market integration only for SP in one case, while inconclusive for RS and MS only in one case too. In all the other cases the result was negative ($d_e > 0.75$), ruling out the possibility of market integration.

Differences in market size and mainly in the proportion of total beef production that is traded in the domestic market derive in differences in price formation between Uruguay and the different Brazilian markets studied in this research. Although Brazil is a key player in the global markets for beef, the size of its domestic market allows buffering the effects of shocks in the international market. In addition, policy issues (taxes, exchange rates) and climate conditions also explain the differences found in the analysis.

4. Conclusions

The importance of market integration has been documented by numerous studies. Specifically, in cattle markets, spatial price relationships have important implications in defining geographic markets, promoting price discovery, and assessing market performance. This study contributes to the cattle market integration literature in two ways. First, it provides a good example of the empirical application of a fractional cointegration approach. Second, No other study in the past addressed market integration in the MERCOSUR cattle market.

Together, the four original members, Brazil, Argentina, Paraguay, and Uruguay, raise 290 million bovine heads (20% of the total world's herd). This is one of the most important beef producing regions on the planet, concentrating about 20% of total world's production, and more than a quarter of global beef exports.

Although the 'separate markets' hypothesis could not be rejected in this particular study, the empirical results denote that market integration among MERCOSUR cattle markets is a plausible approach at some general level, especially when looked with a long run perspective. In theory, the sole existence of an economic bloc should be enough for suspecting the existence of integrated markets.

In the real world, however, things may be a little more complex than that. The MERCOSUR bloc was formally born with the signature of the Asunción Treaty between the four original partners, on March 26, 1991. Twenty-three years have passed since then, with calm and troubled water flowing under the bridge. In many senses, it does not sound unrealistic arguing that the bloc does not behave as a free-trade region within its borders yet. The discussion about the reasons behind the current state of development of the MERCOSUR is well beyond the scope and interest of this study. What matters here only is to see if this configures an advantage or a disadvantage for the integration of the cattle markets within its boundaries.

The basic trends and changes in prices are similar in all cattle markets, as all of them are more or less related to the international markets, within the region and outside the region. The large external shocks, such as the rise and fall of commodity prices of 2008, affected markets all over the planet, and the MERCOSUR markets were not the exception. The global changes in demand and supply, in the preference of consumers, in the industrial organization, of the food industry and more specifically in the meat industry, from production to distribution, from the farm to the table, all of them, have affected cattle markets in these countries to a greater or less extent.

However, there are some dynamics coming from various sources that can separately influence a particular market at a certain point in time, by means of temporary shocks, with different duration and persistence. Some of these factors are market-specific; some are related to particular agro-climatic and ecological conditions under which cattle farming systems are developed, and some are more linked to the micro and macroeconomic conditions prevailing within each country or region.

With regard to the markets, 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 in the different commercial channels, among other things, may introduce some noise in market prices. Some of these factors may derive in product differentiation, hurting the assumption of homogeneity lying under the idea of the law of one price. When comparing price series, some of these market dynamics become apparent most likely in terms of regular gaps between the series rather than in terms of shocks. If these gaps are permanent and more or less keep the same proportion over time, they will not preclude market integration.

Other distortions, such as those caused by climate and environmental factors, including nutritional and sanitary aspects can cause shocks of different entity and persistence, especially when they involve the occurrence of abnormal weather conditions and extreme events (disease outbreaks, droughts, floods). The results obtained in this study showed that in most cases, even when the general behavior of the series under analysis encouraged the idea of some market integration, their response to specific shocks was dissimilar and, most importantly, the reversion of the disturbed series to equilibrium was always very low.

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Table 1. Herd size, beef production, and beef exports of MERCOSUR countries.

Country	Herd Size (heads)	2012 Beef Production		2013 Beef Exports			Markets
		TMT[(1)]	%	TMT[(1)]	%	Million US$	
Brazil	217.4	9,307	73.5	1,184	66.4	5,359	98
Argentina	51.0	2,500	19.7	152	8.5	1,293	59
Paraguay	13.4	355	2.8	213	11.9	991	38
Uruguay	11.5	500	3.9	236	13.2	1,177	58
Total	**293.3**	**12,662**	**100.0**	**1,785**	**100.0**	**8,820**	**120**

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

Source: Based on FAOSTAT, 2014 (herd size and production) and URUNET, 2014 (beef exports and destination markets). All data correspond to year 2013 except production data (year 2012).

Table 2. Cointegration analysis between cattle prices of Uruguay (UY) and São Paulo (SP).

<table>
<tr><th rowspan="2">Value of m</th><th colspan="3">Order of Integration</th><th rowspan="2">Cointegration Coefficient</th></tr>
<tr><th>UY price</th><th>RS price</th><th>Error</th></tr>
<tr><td rowspan="4">$m = 20$</td><td>$d_y = 0.774$</td><td>$d_x = 1.040$</td><td>$d_e = 0.485$</td><td rowspan="3">$\beta = 0.663$</td></tr>
<tr><td>CI: (0.582; 0.967)</td><td>CI: (0.889; 1.190)</td><td>CI: (0.342; 0.629)</td></tr>
<tr><td>p-value: 0.0008</td><td>p-value: 0.0000</td><td>p-value: 0.0034</td></tr>
<tr><td colspan="4">Yes. Evidence of fractional cointegration. d_y not different from d_x and $d_e < 0.5$</td></tr>
<tr><td rowspan="4">$m = 30$</td><td>$d_y = 0.971$</td><td>$d_x = 1.020$</td><td>$d_e = 0.921$</td><td rowspan="3">$\beta = 0.663$</td></tr>
<tr><td>CI: (0.851; 1.090)</td><td>CI: (0.931; 1.108)</td><td>CI: (0.805; 1.038)</td></tr>
<tr><td>p-value: 0.0000</td><td>p-value: 0.0000</td><td>p-value: 0.0000</td></tr>
<tr><td colspan="4">No. The error is non stationary ($d_e > 0.5$)</td></tr>
<tr><td rowspan="4">$m = 40$</td><td>$d_y = 1.033$</td><td>$d_x = 1.075$</td><td>$d_e = 0.876$</td><td rowspan="3">$\beta = 0.663$</td></tr>
<tr><td>CI: (0.968; 1.098)</td><td>CI: (1.151; 1.151)</td><td>CI: (0.791; 0.961)</td></tr>
<tr><td>p-value: 0.0000</td><td>p-value: 0.0000</td><td>p-value: 0.0000</td></tr>
<tr><td colspan="4">No. The error is non stationary ($d_e > 0.5$)</td></tr>
</table>

Number of observations: 501

Table 3. Cointegration analysis between cattle prices of Uruguay (UY) and Rio Grande do Sul (RS).

Value of *m*	**Order of Integration**			**Cointegration Coefficient**
	UY price	**RS price**	**Error**	
$m = 20$	$d_y = 0.774$	$d_x = 0.960$	$d_e = 0.533$	$\beta = 0.607$
	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	
	Not conclusive. Slow adjustment to equilibrium after shocks as $0.5 < d_e < 0.75$			
$m = 30$	$d_y = 0.971$	$d_x = 1.03961$	$d_e = 0.908$	$\beta = 0.\ 607$
	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 ($d_e > 0.5$)			
$m = 40$	$d_y = 1.033$	$d_x = 1.067$	$d_e = 0.904$	$\beta = 0.\ 607$
	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 ($d_e > 0.5$)			

Number of observations: 501

Table 4. Cointegration analysis between cattle prices of Uruguay (UY) and Mato Grosso do Sul (MS).

Value of *m*	**Order of Integration**			**Cointegration Coefficient**
	UY price	**MS price**	**Error**	
$m = 20$	$d_y = 0.774$	$d_x = 1.050$	$d_e = 0.526$	$\beta = 0.636$
	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	
	Not conclusive. Slow adjustment to equilibrium after shocks as $0.5 < d_e < 0.75$			
$m = 30$	$d_y = 0.971$	$d_x = 0.981$	$d_e = 0.912$	$\beta = 0.\ 636$
	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 non stationary ($d_e > 0.5$)			
$m = 40$	$d_y = 1.033$	$d_x = 1.047$	$d_e = 0.902$	$\beta = 0.\ 636$
	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 non stationary ($d_e > 0.5$)			

Number of observations: 501

Table 5. Cointegration analysis between cattle prices of Uruguay (UY) and Argentina (AR).

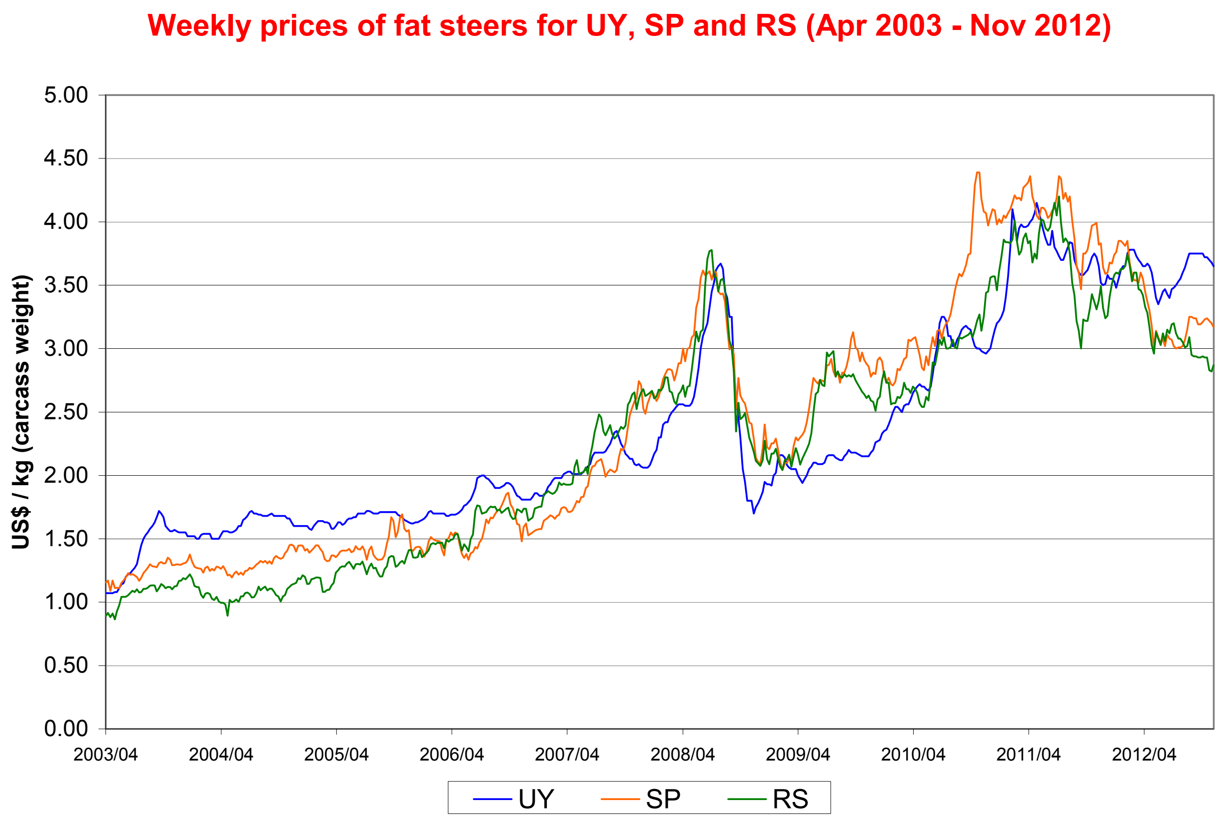
Value of *m*	**Order of Integration**			**Cointegration Coefficient**
	UY price	**AR price**	**Error**	
$m = 20$	$d_y = 0.774$	$d_x = 1.218$	$d_e = 0.550$	$\beta = 0.448$
	CI: (0.582; 0.967)	CI: (1.045; 1.392)	CI: (0.385; 0.716)	
	p-value: 0.0008	*p*-value: 0.0000	*p*-value: 0.0037	
	Not conclusive. Slow adjustment to equilibrium after shocks as $0.5 < d_e < 0.75$			
$m = 30$	$d_y = 0.971$	$d_x = 1.089$	$d_e = 0.889$	$\beta = 0.448$
	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 ($d_e > 0.5$)			
$m = 40$	$d_y = 1.033$	$d_x = 1.045$	$d_e = 0.949$	$\beta = 0.448$
	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 ($d_e > 0.5$)			

Number of observations: 501

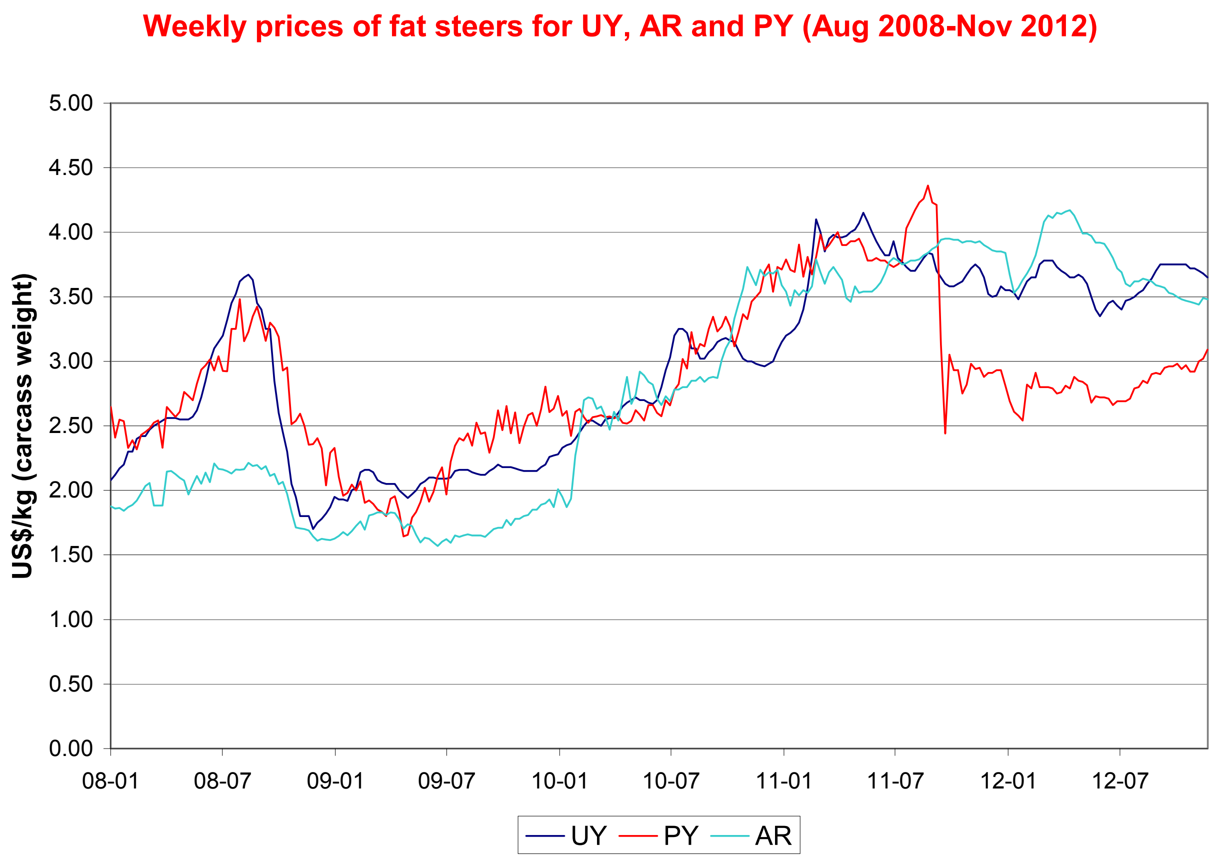
Table 6. Cointegration analysis between cattle prices of Uruguay (UY) and Paraguay (PY).

<table>
<tr><th rowspan="2">Value of m</th><th colspan="3">Order of Integration</th><th rowspan="2">Cointegration Coefficient</th></tr>
<tr><th>UY price</th><th>PY price</th><th>Error</th></tr>
<tr><td rowspan="4">$m = 20$</td><td>$d_y = 0.922$</td><td>$d_x = 1.242$</td><td>$d_e = 0.538$</td><td rowspan="3">$\beta = 0.927$</td></tr>
<tr><td>CI: (0.841; 1.004)</td><td>CI: (1.053; 1.430)</td><td>CI: (0.526; 0.550)</td></tr>
<tr><td>p-value: 0.0000</td><td>p-value: 0.0000</td><td>p-value: 0.0164</td></tr>
<tr><td colspan="4">Not conclusive. Slow adjustment to equilibrium after shocks as $0.5 < d_e < 0.75$</td></tr>
<tr><td rowspan="4">$m = 30$</td><td>$d_y = 1.079$</td><td>$d_x = 1.038$</td><td>$d_e = 0.737$</td><td rowspan="3">$\beta = 0.927$</td></tr>
<tr><td>CI: (1.012; 1.146)</td><td>CI: (0.911; 1.165)</td><td>CI: (0.710; 0.764)</td></tr>
<tr><td>p-value: 0.0000</td><td>p-value: 0.0000</td><td>p-value: 0.0000</td></tr>
<tr><td colspan="4">Not conclusive. Slow adjustment to equilibrium after shocks as $0.5 < d_e < 0.75$</td></tr>
<tr><td rowspan="4">$m = 40$</td><td>$d_y = 1.109$</td><td>$d_x = 1.123$</td><td>$d_e = 0.933$</td><td rowspan="3">$\beta = 0.927$</td></tr>
<tr><td>CI: (1.097; 1.121)</td><td>CI: (1.099; 1.147)</td><td>CI: (0.912; 0.954)</td></tr>
<tr><td>p-value: 0.0000</td><td>p-value: 0.0000</td><td>p-value: 0.0000</td></tr>
<tr><td colspan="4">No. The error is non stationary ($d_e > 0.75$)</td></tr>
</table>

Number of observations: 421



Graph 1. Weekly prices of fat steers for UY, SP, and RS (Apr 2003-Nov-2012).



Graph 2. Weekly prices of fat steers for UY, AR, and PY (Aug 2008-Nov-2012).