MARKET INTEGRATION AND PRICING EFFICIENCY, EMPIRICAL ANALYSES TO THE AGRIBUSINESS SECTOR

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Contents

List of Figures...........................................................................................................vii

List of Tables............................................................................................................. vii

Chapter 1: Introduction ............................................................................................. 1

Chapter 2: Measuring market transaction costs....................................................... 4

2.1. Transportation, transaction and exchange costs............................................4
2.2. The cointegration analysis..............................................................................5

Chapter 3: Transaction costs and trade liberalization: An empirical perspective from
the MERCOSUR agreement .......................................................................................10

3.1. Introduction....................................................................................................12
3.2. Literature review..........................................................................................14
3.3. The relevance of transaction costs in international trade .............................16
3.4. Estimation strategy.......................................................................................18
3.5. Results and Discussion.................................................................................21
3.6. Concluding remarks....................................................................................29
3.7. Bibliographic References...............................................................................31

Chapter 4: Driving factors of agribusiness stock markets: a panel cointegration
analysis.....................................................................................................................37

4.1. Introduction....................................................................................................39
4.2. Drivers of regional stock market integration.................................................41
4.3. Methodology..................................................................................................43
4.4. Data description............................................................................................45
4.5. Empirical results and discussion.................................................................48
4.6. Conclusion.................................................................................................56
4.7. References.................................................................................................57

Chapter 5: Spatial market integration and fuel prices: an empirical analysis of the
Chilean horticultural sector..............................................................................61

3.1. Introduction..................................................................................................63
3.2. The spatial market integration .................................................................64
3.3. Methods .....................................................................................................66
3.4. Data description..........................................................................................68
3.5. Results .......................................................................................................69
3.6. Discussion....................................................................................................73
3.7. Conclusions ...............................................................................................75
3.8. Bibliographic References ..........................................................................76
List of Figures

Figure 1-3. Values of the Gonzalo-Pitarakis sup LMy Statistics with respect to the Chilean Fuel Prices .................................................................70

List of Tables

Table 1-1. Description of the variables and their MFN tariffs and market share in the Brazilian market................................................................. 21
Table 2-1. Results test on cointegration and lag length for each country pairs according to the ADF Residual, Johansen-Juselius tests and Akaike Criterion respectively.......... 22
Table 3-1. TVECM Estimation Results for the pre- and post MERCOSUR periods: Brazil-Argentina................................................................. 24
Table 4-1. TVECM Estimation Results for the pre- and post MERCOSUR periods: Brazil-United States................................................................. 25
Table 1-2. List of trade blocs, member countries and their respective stock markets......................................................................................... 46
Table 2-2. Description of the covariates in the stock market integration model.............................................................................................. 47
Table 3-2. Summary statistics and unit root test for the panel variables ............48
Table 4-2. Correlation matrix for Panel Series of the Model..........................49
Table 5-2. Hausman and F tests for model selection of the panel regression........50
Table 6-2. Estimation results for the models A and B......................................51
Table 1-3. Description of the variables considered in this study......................68
Table 2.3. Estimates for the regime dependent VECM...................................72
CHAPTER 1.

Introduction

Markets are organizations created to facilitate exchange. This is possible because a market comprises a set of institutions developed primarily to reduce a particular type of transaction costs: the market transaction costs. Those costs are important not assess market efficiency, but also to understand how and in which circumstances two or more markets are interconnected. They are the key-variable to understand efficiency in economic organizations.

Transaction costs literature has been widely applied in many different research fields, such as development, finance, agricultural economics and natural resource economics, among others. Despite this recognition, measuring transaction costs is still a challenging task. As Allen (2006) states, if transaction costs could be measured with reasonable accuracy, the theory would become more valuable. More integrated markets are associated with a higher degree of relationship-specific assets, prices or more frequent exchange.

Among the empirical methods applied, Richman and Macher (2006) indicate the main strategies to analyze transaction cost, namely: (i) the qualitative case studies, (ii) quantitative single industry studies, and (iii) econometric analyses. In the econometric analyses, two methods were preferred: (a) time series analysis; and (b) panel data estimation. The advantage of the first method is the possibility of correcting the selection bias associated with estimating the effect of organizational mode on performance (Masten, 1993). The panel data models are useful because they offer many procedures to control unobservable components.

There is an abounding literature exploring the degree of integration between markets and trying to measure market transaction costs, among which the cointegration analyses is one of the most popular. One important shortcoming of this literature is the entanglement of exchange, transportation and transaction costs, due to the inaccuracy about how these concepts are defined and quantified. This imprecision has induced the misleading assumption of transaction costs and, consequently, has yielded an inaccurate measurement of the latter. Actually, the lack of a standardized transaction costs definition is a
shortcoming of many attempts in different research fields that have already tried to quantify them. In order to overcome this issue and by means of cointegration models, this thesis suggests a different set of procedures to measure market transaction costs in different agricultural sectors and can be easily replicated to different markets. This procedure also allows to measuring the exchange and transportation costs associated with a transaction and the effect transaction cost has on the agricultural sector.

This PhD thesis consists of three papers which explore price transmission and market efficiency in selected fields of agricultural economics.

The first, investigate whether Brazil became more integrated with reduced transaction costs after the introduction of MERCOSUR with respect to its main agricultural trade partners, Argentina (a MERCOSUR member) and the United States (a non-MERCOSUR member). Using a threshold vector error correction model (TVECM), we estimate the transaction cost, price transmission elasticity and half-lives adjustments for the most traded agricultural products between Brazil/Argentina and Brazil/United States from January 1980 to December 2012.

The second, explores the drivers of regional stock market integration with a focus on the agribusiness sector across the most important regional trade blocs around the world. Based on the literature on market integration and stock return pricing, we identify nine possible determinants of stock market integration, which we separate into three categories: individual market performance, macroeconomic conditions and agricultural trade. We implement panel cointegration models to analyze the stock indices of agribusiness firms in MERCOSUR, EU, APEC and NAFTA. Furthermore, we account for agriculture-specific factors to control for possible structural shifts in financial markets by including the two main commodity price bubbles during last 20 years.

Finally, the third paper propose a procedure to estimate a regime-dependent vector error correction model with an exogenous threshold variable (Chilean retail fuel prices) were not only the short and long-run equilibrium relationship itself can display threshold-type non-linearity. The proposed approach is unique in explicitly testing the threshold cointegration process based on the Gonzalo and Pitarakis (2006) test. We considered the most relevant central and regional wholesale markets (Santiago and Talca) the prices of the most planted Chilean horticultural products, namely: maize, tomato, onion, carrot. In order to account for
the fuel prices, we consider the Chilean retail prices. The research was conducted using a price series of weekly frequency for the period January 2009 to December 2013.
CHAPTER 2.

Measuring market transaction costs

2.1. Transportation, transaction, and exchange costs

According to Serigati and De Azevedo (2014), exchanging a good is a costly activity. As well as it is necessary to incur expenses to produce a good, it is also necessary to allocate resources to market it. According to Coase (1988), market is an institution that exists “to facilitate exchange”, that is, it exists “in order to reduce the cost of carrying out exchange transactions”. Thus, we assume for now on that market is an organization created to facilitate the exchange. It is possible because a market is composed of a set of institutions that reduce the market transaction costs. Those costs are associated with the necessary activities to elaborate and to enforce the contract that will intermediate the exchange.

The literature that has tried to measure market transaction costs frequently has considered those costs synonymous to transportation costs. However, transportation and transaction costs represent costs of different origins; while transportation costs represent the costs of transfer physically a good from one market to another, taking into account fuel, freight, taxes, tariffs, wages, fares, etc., market transaction costs are linked essentially to information and bargaining costs. Moreover, combined the market transaction costs and the transportation costs form the exchange costs. This idea is reflected in Benham and Benham (2001) definition of exchange costs: “the opportunity cost faced by an individual to obtain a specified good using a given form of exchange within a given institutional setting”.

Higher exchange costs make the trading process more expensive. Actually, they can even insulate markets. Markets for the same good can be in equilibrium with different prices because it is costly to ship products from one market to another to take advantage of arbitrage opportunities triggered by a price difference. It is worth to highlight it is not any price difference that can consolidate a profitable arbitrage opportunity; this difference must be higher than (or at least equal to) the exchange costs. Therefore, ceteris paribus, the higher the exchange costs, the lower is the probability of different markets be integrated.

According to Fackler and Goodwin (2001), “market integration is best thought of as a
measure of the degree to which demand and supply shocks arising in one region are transmitted to another region”.

A market reduces exchange costs primarily because it makes the information disclosure process less costly. As information becomes a cheaper asset, become easier to identify possible gains from arbitrage opportunities taking advantage of the price difference in the two or more markets. Thereby better information flow improves the degree of integration between different markets. In sum, both transportation and market transaction costs (i.e. the exchange costs) influence decisively the degree of integration between different markets. Thus, it is necessary to take them into account to analyze empirically the connection of two or more markets.

2.2. The cointegration analysis

Many empirical procedures have been applied to study market integration but the cointegration analysis has been one of the most popular approaches. This literature assumes the degree of price transmission as a proxy for the level of market integration because it can ‘measure’ the market efficiency in taking advantage of possible arbitrage opportunities. Interestingly, Marshal’s (1920) market definition gives support to this assumption: similar goods belong to the same market whenever their prices converge. For the cointegration literature there is convergence when the prices of those goods share the same long-run stochastic trend. This convergence means the existence of a long-run equilibrium influencing the prices behavior in the short run. In this situation, according to Fackler and Goodwin (2001), market integration is usually a measure of the degree of price transmission between different markets and market efficiency is used to denote a situation in which the agents have left no arbitrage opportunities.

There is an important assumption in the linear cointegration models: the prices in the short run adjust to any deviation in the long-run equilibrium, do not matter how small this change is. It is a strong assumption because, as we have already discussed previously, the arbitrage opportunity is profitable only if the price difference is higher than the exchange costs. Meyer and von Cramon-Taubadel (2004) also criticize the assumption of linear and symmetric adjustment in cointegration models. According to them, the asymmetric price
transmissions are more frequently than the symmetric ones due to the presence of: market power, political intervention, inventory management, adjustment costs (menu costs), and asymmetric information (different search costs among the agents involved in the transaction).

Actually, the cointegration literature has already developed non-linear models that incorporate asymmetric adjustments and exchange costs in its analyses\(^1\). Enders and Granger (1988) is one of the first papers to suggest an approach to evaluate price transmission equation with asymmetric price adjustments. Balke and Fomby (1997) suggest a method to incorporate transaction costs in the cointegration models; those models are known as the threshold cointegration models. Briefly, the threshold cointegration models incorporate the transaction costs including nuisance parameters linking the equations of the system. Those nuisance parameters are called thresholds and they allow splitting the system in different regimes. With different regimes, it is possible to evaluate empirically in which situations the prices are linked (i.e. in which situations the markets are integrated), how strong this connection is in each regime, and what the trade flow direction is. Moreover, the value of each threshold is read as a measure of the transaction costs.

Meyer (2004) proposes a procedure to measure indirectly the transaction costs in currency values using the threshold influence on the price transmission equation. With the variables in natural log, the threshold represents how much, in percentage, the deviation has to be above or below the long-term equilibrium to trigger the regime change. This long-term equilibrium is calculated substituting the variables in the cointegration vector by their respective sample mean.

Despite the popularity of the cointegration models, this approach presents several limitations. In the next paragraphs we focus on the misled concept of transaction costs applied in this literature. There are also two conceptual inaccuracies in this literature:

- It has employed the concept of transaction costs improperly: what they have named transaction costs are better expressed by the term exchange costs, the combination of transportation costs and variable market transaction costs;

\(^1\) Asymmetry and exchange costs are not the same thing. The presence of exchange costs can cause asymmetric adjustments, but not all asymmetry is a consequence of exchange costs. However, both can be modeled using the threshold cointegration models.
• Besides using the concept of transaction costs far from the Coase’s idea, this literature use almost only examples of transportation cost to justify the existence of a persistent price difference. Transportation and transaction costs capture different aspects of an exchange; they are not synonymous.

It is possible to cite many examples. Bekkerman et al. (2013) employ a definition of transaction costs close to the idea of exchange costs: “the cost required to transfer a good from one market to another”. However, when they model those transaction costs, they select only variables associated with transportation costs like fuel costs and seasonality components. The same approach is observed in Campenhout (2007) who use transportation variables (steep passes, road bad conditions, heavy traffic, number of police check posts, bribes, and costs of living) as proxies for transaction costs. Perhaps ‘bribes’ can be a reasonable proxy for transaction costs, but his idea about this concept is clearly far from Coase’s idea: “as expected, the estimated transaction costs are generally proportional to the distance between two markets”.

Goodwin et al. (2002) and Stephens et al. (2012) justify a threshold effect in the price transmission because they analyze perishable goods. This is a good proxy for transaction costs; it can be classified as temporal asset specificity. However, Stephens et al. (2012) empirically use only transportation costs variables (fuel prices and bus fares) as proxies for the exchange costs.

There are also papers that justify the existence of transaction costs due to market characteristics, even when those features are not clearly transaction costs. Rapsomanikis and Hallam (2006) suggest that adjustment costs in the sugar-ethanol processing industry and technical factors (the substitution possibilities between ethanol and oil) explain the existence of transaction costs. Among the market characteristics, Park et al. (2002) cite as transaction costs variables that are not all really transaction costs (trade restrictions, infrastructure bottlenecks, managerial incentive reforms, traders skill, market institutions, the lack of influence of future markets, bribes, and productions specialization policies). But at least they recognize that markets do not emerge overnight; it can be a long and costly process.
Many papers, like Serra et al. (2008) and Serra et al. (2006) model possible transaction costs without justifying why can be reasonable to consider them. On the other hand, Balcome et al. (2007) do not present a definition about those costs, but recognize that transportation and transaction costs are expenses with different origins. Moreover, they also recognize the existence of fixed and variable exchange costs.

Nowadays, spatial market integration analysis and the role played by transaction costs has a main role in research and policy making, thus the people conducting such analyses have to be more aware of the theoretical implications in order to address properly the conclusions of their empirical work.

2.3. References


Measuring market transaction cost


CHAPTER 3

Transaction costs and trade liberalization: An empirical perspective from the MERCOSUR agreement

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ABSTRACT

Studies investigating the effect of trade liberalization policies on transaction costs in agricultural markets are scarce. The objective of our paper is to determine whether Brazil became more integrated with reduced transaction costs after the introduction of MERCOSUR with respect to its main agricultural trade partners, Argentina (a MERCOSUR member) and the United States (a non-MERCOSUR member).

Using a threshold vector error correction model (TVECM), we estimate the transaction cost, price transmission elasticity and half-life adjustments for the most traded agricultural products between Brazil/Argentina and Brazil/United States from January 1980 to December 2012. Our findings suggest a strong MERCOSUR effect, with lower transaction costs (TC) and higher price transmission elasticity when compared to a non-agreement scenario. Moreover, the variations of both parameters are highly heterogeneous across products, depending mainly on their degree of differentiation. From a policy perspective, elements such as the sources of comparative and competitive advantages together with investment policies, specific market regulations and agricultural subsidies, among others, are mainly what influence the extent of transaction cost and market integration. Our results show that Brazil has made progress but still has considerable room for improvement in reducing barriers to agricultural products and, as a consequence, to achieving the full benefits of the MERCOSUR agreement.

Keywords: Market Integration, Transaction Costs, Threshold Vector Error Correction Model, Brazil, Argentina, United States.
1. Introduction

Although trade liberalization policies have historically promoted market integration of regions or countries, recent studies suggest that their effectiveness is negatively affected by the presence of transaction costs (TCs) (Listorti, 2009, Stephens et al., 2012). In fact, in order to maximize the benefits from trade liberalization policies, recent studies (Mitra and Josling, 2009; Martin and Anderson, 2009) argue that countries should take actions to identify and reduce sources of TCs between markets. A market integration analysis offers a way to estimate the level of TCs and, consequently, allows for an assessment of whether regional trade liberalization, under different levels of TCs, affects the price transmission degree between countries or regions (Balcombe et al. 2007).

While studies of market integration in a spatially separated contexts have received substantial attention in the literature (e.g. Park et al. [2002] for China; Getnet et al. [2005] for Ethiopia; Cudjoe et al. [2010] for Ghana and Valdes et al. [2011] for Chile), only a few studies have explicitly examined the impact of trade agreements on TCs and their implications for the transmission of price signals between agricultural markets.

The drivers of price transmission include not only the level of trade but also the market determinants for each country (Koester, 2001). The aim of this study is to explore which factors affect the degree of market integration, the routes of TC variations derived from the implementation of regional trade agreements and their effect on the price transmission level between agricultural markets.

In order to accomplish our objective we used as case study the variation on market integration parameters of the Brazilian agricultural market with respect to its major trade partners in both periods: the United States and Argentina, respectively. We used the implementation of the Common Market of the South (MERCOSUR) as a reference of structural break in trade. It is expected that after the agreement Brazilian trade will shift from the United States to Argentina. This study will estimate TCs, price transmission elasticity (PTE) and their implications for market integration on the top nine agricultural products traded from 1980 to 2012 between Brazil and the USA (without any trade agreements) and Argentina and Brazil (existence of a trade agreement).

Previous theoretical studies (see Baulch, 1997 and Blavy and Juvenal, 2009) show that
transaction costs generate a no-trade threshold band where prices in two locations fail to equalize. Outside this threshold band, arbitrage is profitable and trade is promoted, a dynamic that is captured successfully by a threshold vector error correction model (TVECM) (Balke and Fomby, 1997). This main advantage of this model is its ability to analyze the impact of TCs on market integration solely on the basis of price information. In this case, the model is capable of identifying a lower bound of the relative TCs associated with equilibrating price adjustment, e.g., through arbitrage and trade. To the best of our knowledge, this is the first attempt to relate the impact of regional trade agreements on TCs and price transmission parameters in agricultural markets.

MERCOSUR is a customs area implemented in 1995 with Brazil, Argentina, Uruguay, and Paraguay as the original partners. Among these countries, Brazil and Argentina generate more than three-quarters of its agricultural production. Brazil is considered the most significant agricultural market in Latin America and one of the top 10 players in world agricultural trade (GVF, 2013). Before the implementation of MERCOSUR, Brazilian agricultural imports were mainly dominated by the USA, followed by Argentina and the EU (FAOSTAT, 2013). After the implementation of MERCOSUR, the situation changed and Argentina became the number one trade partner, followed by the United States, Venezuela, and China (CONAB, 2013). Even though Brazil has fostered trade openness in order to meet a growing domestic demand for food, there are still signs of high TCs with its main trade partners (Monteiro et al., 2012). Therefore, there is still no clear evidence about the effect of trade agreements with respect to the variation of TCs.

Accordingly, this paper attempts to: first, to perform a comparative analysis to determine if after the implementation of MERCOSUR, TCs between Brazil-Argentina and Brazil-United States were reduced and if this effect implied a higher PTE; and second, to analyze whether the formation of TCs and PTE are product-specific and determined by each market structure.

The article proceeds as follows: Section 2 reviews the Brazilian agricultural market and the characteristics of MERCOSUR. Section 3 gives a glance at the relevant literature on TCs and market integration. Section 4 describes the methodology and data sources. Sections 5 and 6 present the results and discussion, respectively. Finally, Section 7 summarizes the main conclusions.
2. Literature review

2.1 Brazil in the MERCOSUR context

MERCOSUR is an economic and political treaty, whose members are Brazil, Argentina, Uruguay, Paraguay and since 2011 Venezuela, while Bolivia and Chile are associate members\(^2\). This custom area consists in a gradual process of tariff harmonization between these countries with the goal of establishing a common external tariff (CET), which was finally achieved in early 1995 (Bas, 2012). Today, duty-free access is provided to all goods produced within the zone with the exception of automobiles and sugar. Along with the establishment of a CET, the agreement allowed for the free movement of goods, services and production factors, the abolition of restrictions over reciprocal trade, adoption of common trade policies towards countries that do not belong to MERCOSUR and the coordination of macroeconomic and sectorial policies.

In 2012, agriculture accounted for 32% of total member exports, 9% of which corresponding to intra-trade among MERCOSUR countries (FAOSTAT, 2013). According to Korinek and Melatos (2009), this situation could be due to MERCOSUR’s limited effect on developing comparative advantages among its members and the fact that when the CET was established its member economies were engaging trade liberalizations agreements with other markets simultaneously. As a result, trade with non-member countries was not affected and in some cases it even grew.

At a regional level, MERCOSUR’s agricultural market size is largely determined by Brazil (Protil et al., 2010). Regarding Brazilian imports, before the implementation of MERCOSUR the principal provider was the United States (18%), followed by Argentina (17%), the EU (14%), Uruguay (8%) and Paraguay (6%). At the time, the main Brazilian imports were wheat (17%), malt (8%), cotton (6%), potatoes (6%) and agro-industry inputs (4%) (FAOSTAT, 2013). In contrast, after the implementation of MERCOSUR’s, Argentina has become the biggest exporter of agricultural products to Brazil (24%), followed by the United States (18%), Venezuela (14%) and China (17%) (CONAB, 2013).

Currently, Brazil’s imports have continued to be dominated by wheat and its derivatives

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\(^2\) On June 28, 2012, Paraguay was barred from participating in MERCOSUR decisions until it held democratic elections. On July 30, 2012 in Brasilia, the other countries and full members of MERCOSUR approved the final incorporation of Venezuela as a full member, which became effective on August 12, 2012.
(19%), but barley (11%), fresh fish (6%), beans (7%) and fresh pears (7%) have joined the list (FAOSTAT, 2013).

2.2 Main agricultural trade policies of the United States and Argentina

Since the United States is one of the biggest players in the global agricultural market, its main trade policies (namely, market development, export subsidies and market access programs) may have an important effect on price behavior and arbitrage activities in major agricultural markets around the world (Mitra and Josling, 2009). This situation is reinforced by the country’s ample internal logistic network, which allows it to transport agricultural products to international markets cheaply and efficiently (Korinek and Melatos, 2009), allowing the United States to have competitive advantages in access and product shipping, that in turn should result in lower TCs.

Brazil and the United States have a long history in terms of agricultural trade. Before the implementation of MERCOSUR, the United States was the main exporter of agricultural products to Brazil, as mentioned above. In 2010, both countries signed an agreement for trade and economic cooperation, a joint effort to promote mutual trade and investment (Coelho, 2009). As a result, in 2012, Brazil became the seventh largest goods export market for the United States, totaling $1.9 billion (Sumner, 2013). Current leading categories include: wheat (US$1.2 billion), dairy products (US$83 million), prepared food (US$67 million), and feeds and fodders (US$51 million) (CONAB, 2013).

On the other hand, Argentina is the eighth largest producer and the twelfth exporter of agricultural commodities in the world. In 2012 it produced 8.4% of global agricultural output and its products represent, on average, 2.9% of world agricultural trade during the last decade (FAOSTAT, 2013). Nevertheless, when compared to the United States, its transportation and marketing costs for bulk agricultural product exports have historically been much higher (Brum and Kettenhuber, 2008). This is largely due to an inefficient or underdeveloped barge and railroad transportation system and a heavy reliance on more expensive truck hauling that reduces the country’s competitive advantage.

Argentina’s trade policy has historically been one of protectionism, emphasizing import substitution (Bas, 2012). Due to the hyperinflations of 1989 and 1990, the Argentinean
government was forced to shift towards market-oriented policies and launched an ample unilateral trade-liberalization process to promote exports. These policy changes included trade liberalization, deregulation, privatization of many state enterprises, the MERCOSUR implementation and a currency convertibility program\(^3\) which allowed a significant expansion of exports (Sturzenegger and Salazni, 2007). As a result, during 2012, Argentina exported 47% of the world’s soybean oil, 11% of soybeans, 7% percent of wheat and 5% percent of fresh beef (GVF, 2013).

Policies engaged by both Brazilian partners have had a history of considerable differences that target TCs in opposite directions. While the United States encourages trade by reinforcing logistics networks and generates incentive to open markets, Argentina has created barriers along the years that translate into a less favorable trade capacity (Nogues and Porto, 2007). In the case of Argentina, these actions include an export tax, initially at 15%, to restrict exports of meat and dairy products and a complex compensation scheme for wheat and corn to allow domestic users buy these grains at a more favorable price than that available to exporters. These situations clearly affect the level of efficiency in which the tariff and non-tariff advantages of MERCOSUR are implemented between member countries.

3. The relevance of transaction costs in international trade

Price theory literature has provided wide theoretical evidence about the lack of convergence of international prices due to the presence of transaction costs (TCs) (Gonzalez-Rivera and Helfand, 2001; Juvenal and Taylor, 2008). Therefore, not incorporating TCs in the analysis could distort estimations of degree of integration and convergence among markets or regions (Meyer and Von Cramon-Taubadel, 2004). Following the definition given by Barrett (2001), TCs between a market “i” and “j” are composed of transportation costs \((f_{ij})\), where distance is one of the most important factors; variable costs \((v_{ij})\) associated with

\(^3\) The currency convertibility program was designed to eliminate the main source of inflationary pressures, that is, the creation of money to finance the public sector deficit. The convertibility program consisted of a currency board that fixed a nominal relation of one Argentine peso to one U.S. dollar. The currency board was required to provide full backing in U.S. dollars for any issue of Argentine pesos. Moreover, the U.S. dollar was established as legal tender within Argentina.
rates, cargo insurance, contracts, financial expenses, hedging, sanitary and phytosanitary barriers; customs duties (dij); and unmeasurable costs (wij), such as opportunity cost, the cost of searching for information and risk premiums.

Lence and Falk (2005) show that the implementation of trade agreements reduces TCs, generating higher price transmission and therefore a higher integration of the markets. According to these authors, this situation increases the levels of market efficiency and welfare gains mainly due to a closer alignment with the equilibrium condition expressed in the Enke-Samuelson-Takayama-Judge Model (ESTJ). Along the same lines, Barrett (2005) highlights that welfare gains are directly related to the minimization of TCs, because when these are expressive and/or trade barriers are effective, the economy is deprived of the benefits of specialization in trade. His conclusion is supported by Alam et al. (2012), who analyzed market integration between five major rice markets in Bangladesh. Their results highlight the importance of shaping policies to reduce TCs, in order to create greater market efficiency among regional markets.

From an empirical point of view, different components of TCs have been tested for their impact on market integration. For instance, Aker (2008) found that in Niger, where road density and the quality of products are low, the TCs associated with accessing markets tend to be higher for grain than for other markets. Moreover, Pingali (2005) suggests that perishable crops, like vegetables, are usually associated with high TCs when compared to non-perishable products.

The studies that most resemble our work are those of Meyer and Von Cramon-Taubadel (2004) and Amikuzuno (2009). The first studied the effects of TCs on the spatial integration of the pork market between Germany and Holland, concluding that higher TCs results in a lower arbitrage and therefore lower market integration. According to the authors, ignoring TCs in the econometric analysis can lead to erroneous conclusions in terms of the degree of price transmission between countries. The second study analyzed the implications of trade liberalization on market integration between tomato producers and consumer markets during periods of high and low agricultural import tariffs in Ghana. This

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4 The equilibrium concept at the heart of most trade theory is that of Enke-Samuelson-Takayama-Judge (ESTJ): spatial equilibrium, in which the dispersion of prices in two locations for an otherwise identical good is bounded from the top via the cost of arbitrage between the two markets. Trade volumes are unfettered and bounded from below when trade volumes reach a certain ceiling value.

5 Aker (2008) defines road density as the ratio of the length of the country's total road network to the country's land area.
author found that the speeds of price adjustment were higher after import tariffs were lowered. Looking at a specific trade agreement, Blavy and Juvenal (2009) investigated the market integration level in sectorial real exchange rate dynamics between Mexico, Canada and the United States for the periods before and after NAFTA. Their results show that prices adjusted much faster in the post-NAFTA period mainly because of lower TCs between each country pair.

In Brazil, studies have mainly focused on the analysis of market integration parameters for individual sectors or products. For example, González-Rivera and Helfand (2001) analyzed the integration of Brazilian rice markets. The authors demonstrated that distance and quality contribute to the formation of TCs between major Brazilian domestic markets. Coelho (2009) studied the integration of internal and external cotton markets in order to estimate the influence of trade liberalization on price transmission. His results show that the Brazilian and U.S. markets are perfectly integrated. Similar to González-Rivera and Helfand (2001), this author concluded that transport logistics and infrastructure availability directly contribute to increasing market integration through higher levels of price transmission.

Our work fills a gap in the current literature by comparing the effect on TCs and PTE using two country pairs: Brazil-Argentina, with a long history of trade agreements; and Brazil-United States that, on the contrary, although they have a long-term trade relationship do not possess any long-term agreements. The chosen countries make for a valid comparison case study because they rank among the top players in the global agricultural market.

4. Estimation strategy

4.1 Model description

In order to estimate TCs and PTE in this paper we applied a regime-dependent model called Threshold Vector Error Correction Model (TVECM) (Balke and Fomby, 1997). More specifically, we estimated TVECM models for each price pair between Brazil-Argentina and Brazil-United States during the pre- and post-MERCOSUR periods. The main advantages of TVECM are that: a) prior identification of causality between analyzed price series is not necessary; b) the timeline of the price transmission process, and
therefore, of market integration, can be identified; and c) it includes elements that positively influence the price transmission analysis, for example, deterministic trends, stochastic trends (non-stationary) and autocorrelation. We implemented a TVECM with one threshold and two regimes according to the Akaike (AIC) specification criterion. This model can be formulated as follows (1):

\[
\begin{align*}
\Delta p_t & = \begin{cases} 
\rho_1 \Delta p_{t-1} + \theta_1 + \sum_{m=2}^{M} \theta_{1,m} \Delta p_{t-m} + \varepsilon_t & \text{if } \gamma' p_{t-1} \leq \psi \\
\rho_2 \Delta p_{t-1} + \theta_2 + \sum_{m=2}^{M} \theta_{2,m} \Delta p_{t-m} + \varepsilon_t & \text{if } \psi < \gamma' p_{t-1} 
\end{cases} \\
& \text{(Regime 1)} \\
& \text{(Regime 2)}
\end{align*}
\]

where \( p_t = (p_{t,1}, p_{t,2}) \) is the corresponding observation in the period \( t=1\ldots n \), of a two dimensional time series generated by a TVECM with 2 regimes, which are characterized by the parameters \( \rho, \theta_k \) and \( \theta_{km} \) for \( k=1,2 \) and \( m=1,\ldots,M \). We call "\( \psi \)" the threshold parameter that allows for an asymmetric price adjustment to the long run equilibrium. It can be interpreted as the TCs for moving a product from one market to another (Ihle and Von Cramon-Taubadel, 2008). As Baulch (1997) states, this parameter represents the band within which trade between both markets or sectors would not be profitable.6

A critical element that might cause a bias in the estimation of each TVECM is an imprecise estimation of TCs and their transition between each regime. Currently, most threshold estimators are based on a profile likelihood function, which are especially prone to be unreliable in situations characterized by large numbers of unknown model parameters besides the thresholds. This occurs when there is little difference between adjoining regimes, and when the location of the estimated thresholds leaves only few observations in one of the regimes (Balcombe et al. (2007)).

In order to overcome this issue, we estimated the threshold parameter by using a regularized Bayesian estimator with a posterior density (Greb et al., 2013). This offers an accurate estimation of the TCs and produces results that are more consistent with the theory of spatial equilibrium than the corresponding profile likelihood results.

---

6According to Balcombe et al. (2007), in the case of a threshold parameter of, for example, 0.15, both positive and negative price variations of up to 15% from the equilibrium price are acceptable to traders. Only if prices diverge more than 15% from equilibrium will arbitrage activity be triggered. In this case, the threshold band is estimated at 30% (15% above and 15% below equilibrium).
Market integration was analyzed by the traditional price transmission elasticity (PTE) and its corresponding half-life coefficient. While the first is the long run equilibrium parameter ($\gamma$), the second depends on the adjustment speed of the model’s outer regime ($\rho_{ke}$) and is defined as the time required for the effect of 50% of a price shock to phase out. It is calculated by the equation: $\ln(0.5)/\ln(\rho_{ke})$.

Prior to the TVECM estimation, we conducted the Elliot-Rothenberg-Stock (ERS) and Kwiatkowski–Phillips–Schmidt–Shin (KPSS) tests for non-stationarity. In addition, we tested for the presence of cointegrating vector(s) for all price series by using the Johansen-Juselius and ADF Residual test procedures simultaneously. We also conducted the Hansen and Seo test (2002) of linear vs. threshold cointegration. We tested heterogeneity, autocorrelation and non-normality with the Alexandersson SNHT, Breusch-Godfrey LM and Lomnicki-Jarque-Bera tests, respectively.

4.2. Data

We used the monthly domestic Brazilian and international FOB prices (in U.S. dollars/ton) for the top nine agricultural products traded between Brazil-Argentina and Brazil-United States from January 1980 to December 2012, resulting in 6244 observations.

Domestic Brazilian prices were converted from Brazilian Reals to U.S. Dollars using free exchange rates (GVF, 2013). All prices series were converted to natural logarithm prior to estimation and testing. Additionally, previous authors suggest that wheat quality, expressed by the protein content, plays an important role in the arbitrage mechanism between milling industries (Brum and Kettenhuber, 2008). We therefore separated wheat price series according to the protein content of the varieties from Argentina and the United States, that is, 11% and 12% for PAN (in the case of Argentina) and 11% and 15% for Hard Red Winter and Soft Red Winter (in the case of the United States).

The data was obtained from the Brazilian National Supply Company (CONAB), the Economic Institute of the State of Sao Paulo (IESP-Brazil), the United States Department of Agriculture (USDA), the U.S. Wheat Associates (USW) and the Argentine Association of Wheat Producers (AAPOTRIGO). Table 1 shows a description of the products included in this study, together with their respective market share and Most Favored Nation (MFN).
tariffs in the Brazilian market.

Table 1. Description of the variables and their MFN tariffs and market share in the Brazilian market

<table>
<thead>
<tr>
<th>Trade Partner</th>
<th>HS Code</th>
<th>Description</th>
<th>Variable</th>
<th>1980-2012 Average % Share Total Agricultural Imports</th>
<th>2012 MFN tariff (%)</th>
<th>2012 Country Share on this HS Code (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>ARGENTINA</td>
<td>100190</td>
<td>PAN 10% protein</td>
<td>Pan10</td>
<td>41.5</td>
<td>0</td>
<td>58</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>100190</td>
<td>PAN 12% protein</td>
<td>Pan12</td>
<td>8.5</td>
<td>12</td>
<td>87</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>110100</td>
<td>Wheat Flour</td>
<td>Whf</td>
<td>6.1</td>
<td>14</td>
<td>35</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>110710</td>
<td>Malt</td>
<td>Mlt</td>
<td>3.1</td>
<td>10</td>
<td>39</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>7133</td>
<td>Beans (Kidney and White)</td>
<td>Kwb</td>
<td>2.9</td>
<td>10</td>
<td>72</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>80820</td>
<td>Pears Fresh</td>
<td>Pef</td>
<td>2.5</td>
<td>14</td>
<td>65</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>200570</td>
<td>Olives</td>
<td>Oli</td>
<td>2.2</td>
<td>7</td>
<td>85</td>
</tr>
<tr>
<td>ARGENTINA</td>
<td>100300</td>
<td>Barley</td>
<td>Bar</td>
<td>1.9</td>
<td>14</td>
<td>74</td>
</tr>
<tr>
<td>USA</td>
<td>100190</td>
<td>HRW 11% protein</td>
<td>Hrw11</td>
<td>41.5</td>
<td>10</td>
<td>21</td>
</tr>
<tr>
<td>USA</td>
<td>100190</td>
<td>SRW 15% protein</td>
<td>Srw15</td>
<td>6.2</td>
<td>13</td>
<td>36</td>
</tr>
<tr>
<td>USA</td>
<td>210690</td>
<td>Food Preparations</td>
<td>Fop</td>
<td>4.9</td>
<td>8</td>
<td>66</td>
</tr>
<tr>
<td>USA</td>
<td>200990</td>
<td>Vegetable Juices</td>
<td>Vgj</td>
<td>3.3</td>
<td>8</td>
<td>16</td>
</tr>
<tr>
<td>USA</td>
<td>382370</td>
<td>Industrial Alcohols</td>
<td>Ina</td>
<td>2.0</td>
<td>14</td>
<td>94</td>
</tr>
<tr>
<td>USA</td>
<td>200520</td>
<td>Potatoes</td>
<td>Pot</td>
<td>1.7</td>
<td>2</td>
<td>13</td>
</tr>
<tr>
<td>USA</td>
<td>100300</td>
<td>Barleys</td>
<td>Bar</td>
<td>1.5</td>
<td>10</td>
<td>44</td>
</tr>
</tbody>
</table>

Source: Prepared by the authors with information from Brazilian Customs Service (2012) and FAOSTAT (2013)
Note: All price series were converted to natural logs prior to estimation and testing.

5. Results and Discussion

5.1 Preliminary tests

According to the KPSS and ERS tests results, all series are I (1). Moreover, as can be seen in Table 2, the Johansen-Juselius and/or ADF Residual tests show that both price pairs are cointegrated during the pre- and post- MERCOSUR periods.
Table 2. Results test on cointegration and lag length for each country pairs according to the ADF Residual, Johansen-Juselius tests and Akaike Criterion respectively.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Pre-MERCOSUR</th>
<th>Post-MERCOSUR</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>ADF Residual Based Test</td>
<td>Johansen-Juselius Trace Test, r=0</td>
</tr>
<tr>
<td>Brazil-Argentina</td>
<td>Pan10 (-) 4.012*** 19.50 **</td>
<td>4 (-) 2.001*</td>
</tr>
<tr>
<td></td>
<td>Pan12 (-) 2.165* 18.90 *</td>
<td>4 (-) 2.221*</td>
</tr>
<tr>
<td></td>
<td>Whf (-) 2.356* 31.44***</td>
<td>2 (-) 3.331**</td>
</tr>
<tr>
<td></td>
<td>Mlt (-) 5.115** 21.44**</td>
<td>2 (-) 7.556**</td>
</tr>
<tr>
<td></td>
<td>Kw (r) 6.511*** 22.35**</td>
<td>2 (-) 4.331**</td>
</tr>
<tr>
<td></td>
<td>Pf (-) 2.569* 17.33*</td>
<td>3 (-) 3.761***</td>
</tr>
<tr>
<td></td>
<td>Oli (-) 7.541*** 22.55**</td>
<td>1 (-) 5.233***</td>
</tr>
<tr>
<td></td>
<td>Bar (-) 3.892** 15.55*</td>
<td>1 (-) 4.213**</td>
</tr>
<tr>
<td></td>
<td>Mlp (-) 3.949** 16.01*</td>
<td>2 (-) 2.188*</td>
</tr>
<tr>
<td>Brazil-United States</td>
<td>Hrw11 (-) 3.790*** 15.20 *</td>
<td>2 (-) 2.256*</td>
</tr>
<tr>
<td></td>
<td>Srw13 (-) 6.015*** 31.08 ***</td>
<td>4 (-) 3.087***</td>
</tr>
<tr>
<td></td>
<td>Fp (-) 1.077 13.22 *</td>
<td>4 (-) 2.141*</td>
</tr>
<tr>
<td></td>
<td>Cot (-) 3.790*** 18.55 **</td>
<td>4 (-) 5.610***</td>
</tr>
<tr>
<td></td>
<td>Anf (-) 2.998* 36.99 ***</td>
<td>4 (-) 3.755***</td>
</tr>
<tr>
<td></td>
<td>Vgj (-) 3.265** 21.45 **</td>
<td>2 (-) 4.647***</td>
</tr>
<tr>
<td></td>
<td>Ina (-) 4.767*** 27.86 ***</td>
<td>3 (-) 2.183*</td>
</tr>
<tr>
<td></td>
<td>Pot (-) 3.681** 16.03 **</td>
<td>3 (-) 4.771***</td>
</tr>
<tr>
<td></td>
<td>Bar (-) 4.888*** 37.66 ***</td>
<td>2 (-) 5.782***</td>
</tr>
</tbody>
</table>

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

In order to test whether the non-linear specification is superior, we used the Hansen and Seo test. This test shows statistical significance favoring a non-linear specification. Ultimately, the residuals did not present heterogeneity and autocorrelation. All test results are available upon request.

5.2. Model results and discussion
5.2.1 Transaction Costs

Table 3 and 4 shows evidence of a strong MERCOSUR effect for the Brazil-Argentina and Brazil-United States pairs. It is worth noting that the reduction in TCs occurs both in Argentina, a member of MERCOSUR, and in the United States, which is not a member of this treaty. Moreover, it is important to notice that TCs formation appears to be product-
specific and highly heterogeneous. Our findings also suggest that the heterogeneity in TCs among products also translate into a heterogeneous change in TCs by each product in the list, a finding previously reported in the work of Blavy and Juvenal (2009). When comparing absolute values during both periods, the Brazil–United States pair presented lower threshold values than Brazil–Argentina. One explanation is derived from the policy strategies implemented in both countries, as described in section 2.2. In particular, the effect of a more open trade policy from the United States during the last 30 years specially facilitates the logistic infrastructure for agricultural imports and exports, which resulted in lower absolute TCs compared to Argentina.
Table 3. TVECM Estimation Results for the pre- and post MERCOSUR periods: Brazil-Argentina

<table>
<thead>
<tr>
<th>Variables</th>
<th>Pre-MERCOSUR</th>
<th>Post-MERCOSUR</th>
<th>Difference in %</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Hansen &amp; Seo test</td>
<td>Half Lives</td>
<td>Half Lives</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Regime 1 (months)</td>
<td>Regime 2 (months)</td>
</tr>
<tr>
<td>PAN 10% protein</td>
<td>0.00</td>
<td>2.441</td>
<td>1.988</td>
</tr>
<tr>
<td>PAN 12% protein</td>
<td>0.00</td>
<td>2.761</td>
<td>2.144</td>
</tr>
<tr>
<td>Wheat Flour</td>
<td>0.01</td>
<td>1.976</td>
<td>1.344</td>
</tr>
<tr>
<td>Malt</td>
<td>0.00</td>
<td>1.541</td>
<td>0.998</td>
</tr>
<tr>
<td>Beans</td>
<td>0.05</td>
<td>2.451</td>
<td>2.004</td>
</tr>
<tr>
<td>Pears Fresh</td>
<td>0.03</td>
<td>3.251</td>
<td>2.664</td>
</tr>
<tr>
<td>Olives</td>
<td>0.08</td>
<td>2.515</td>
<td>2.114</td>
</tr>
<tr>
<td>Barley</td>
<td>0.00</td>
<td>2.033</td>
<td>1.558</td>
</tr>
<tr>
<td>Milk Powders</td>
<td>0.07</td>
<td>1.987</td>
<td>1.225</td>
</tr>
</tbody>
</table>

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively
Table 4. TVECM Estimation Results for the pre- and post MERCOSUR periods: Brazil-United States

<table>
<thead>
<tr>
<th>Variables</th>
<th>Pre-MERCOSUR</th>
<th></th>
<th>Post-MERCOSUR</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>p-value</td>
<td>Regime 1 (months)</td>
<td>Regime 2 (months)</td>
<td>(Ψ₁)</td>
</tr>
<tr>
<td>HRW 11% protein</td>
<td>0.01</td>
<td>2.009</td>
<td>1.224</td>
<td>0.165</td>
</tr>
<tr>
<td>SRW 15% protein</td>
<td>0.10</td>
<td>3.960</td>
<td>2.445</td>
<td>0.415</td>
</tr>
<tr>
<td>Food Preparations</td>
<td>0.03</td>
<td>1.567</td>
<td>1.114</td>
<td>0.310</td>
</tr>
<tr>
<td>Cotton</td>
<td>0.00</td>
<td>2.492</td>
<td>2.105</td>
<td>0.171</td>
</tr>
<tr>
<td>Animal Feed</td>
<td>0.09</td>
<td>3.306</td>
<td>2.441</td>
<td>0.229</td>
</tr>
<tr>
<td>Vegetable Juices</td>
<td>0.12</td>
<td>2.557</td>
<td>2.105</td>
<td>0.137</td>
</tr>
<tr>
<td>Industrial Alcohols</td>
<td>0.10</td>
<td>1.907</td>
<td>1.110</td>
<td>0.218</td>
</tr>
<tr>
<td>Potatoes</td>
<td>0.00</td>
<td>2.067</td>
<td>1.215</td>
<td>0.110</td>
</tr>
<tr>
<td>Barleys</td>
<td>0.03</td>
<td>1.118</td>
<td>0.554</td>
<td>0.287</td>
</tr>
</tbody>
</table>

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively
Taking Barrett’s definition (2001) as a reference, the decrease on threshold values could be a consequence of a reduction in the formative components of TCs after MERCOSUR. These components, namely non-tariff barriers, customs procedures, certain phytosanitary measures, transport and storage times may have all been reduced or relaxed, thus promoting greater or faster access to the Brazilian market.

Taking a further look at sectorial characteristics, however, it is difficult to establish that the implementation of MERCOSUR resulted in a constant development of competitive advantages between Argentina and Brazil. Relative inefficiencies in logistic aspects, due to the lack of permanent investment policies in infrastructure, have affected the level of efficiency in the supply chain for most of the agricultural products traded between both countries\(^7\). This situation clearly affects the extent to which those countries take advantage of the tariff and non-tariff benefits of MERCOSUR.

It is worth noticing that after the treaty the range between the highest and lowest TCs decreased. For the Brazil-Argentina pair during the pre-MERCOSUR period, threshold values ranged from 11% (olives) to 81% (PAN wheat 11% protein), while after the agreement came into force, TCs were reduced to a range from 8.1% (olives) to 44.5% (PAN wheat 10% protein), 5.5 times the value.

On the other hand, for the Brazil-United States pair, after the MERCOSUR, the biggest reductions were for high protein HRW and SRW wheat, industrial alcohol (a product traded in a sector subject to high levels of taxation) and cotton (a highly subsidized producer sector) (Ben-kaabia et al. 2005; Kupfer, 2011).

During this period, the average reduction in TCs in the case of Brazil-United States was 10% when compared to Brazil-Argentina, with a 26% reduction. These results could be used as evidence to conclude that the agreement did have a real impact on TCs, and could explain the shift in Brazilian imports from United States to Argentina after the MERCOSUR came into force.

Now, if we examine the pathways of TCs reduction between a member (Brazil) and a non-member (USA) of MERCOSUR after 1995. One approach is to analyze the terms of trade between Brazil and the United States during post- MERCOSUR period. Previous authors (Maggian and Felipe, 2009; Donoso et al. 2011) confirmed that MERCOSUR created

\(^7\) For a detailed discussion of this issue, see Maggian and Felipe (2009), Kupfer (2011) and reference therein.
increased investment and trade opportunities in Latin America. In particular, Coelho (2009) reported that exports of differentiated agricultural products from non-members to Brazil increased significantly from 1995. Since Brazil and Argentina share similar comparative advantages in a wide group of agricultural products and the structure of comparative and competitive advantages among Brazil, Argentina and United States are highly asymmetric, thus it is not surprising a reduction pattern of TCs between Brazil and the United States after 1995. This situation happens mainly because Argentina and Brazil could avoid trading with each other in agricultural products in which their comparative advantage structure is similar, allowing to open or maintain the market share of US agricultural products in Brazilian market. Within this context, since the TCs formation is product specific, the arbitrage between Brazil and the United States could be promoted by the competitive position of the United States on products subject to price differentiation or product specificity. For example, considering the importance of wheat quality on its price formation and according our model's results, we infer that Brazilian importers chose the US varieties over Argentinean products independently of the existence of MERCOSUR, because of the higher protein content of HRW and SRW compared to PAN varieties (Brum and Kettenhuber, 2008), thus driving more intense exports to important flour producer, such as Brazil.

5.2.2. Price transmission elasticity and half-life coefficients

As described in the methodology section, the TVECM captures the market integration pattern through a joint estimation of TCs and PTE coefficients. Additionally, from the price adjustment parameters, we estimates the half-life adjustments to shocks, which represents the speed at which shocks the variables respond to return the long run equilibrium. Comparing both country pairs before the implementation of MERCOSUR (Tables 3 and 4), it can be seen that the average PTE for Brazil-Argentina is lower than for the Brazil-United States pair. This result suggests that, even though Argentina is closer geographically, the United States leads the international prices for a diverse group of agricultural products and thus generates a higher integration among countries.

As in the case of TC, PTE and half-life are product-specific, with higher PTEs for more differentiated products such as milk powder, industrial alcohol and/or vegetables juices. Half-
life estimates suggest faster adjustments among these types of products. This is a standard result in the literature, reported in studies for other country pairs (see Juvenal and Taylor, 2008).

Similarly to TCs dynamics, PTE and half-life coefficients suggest a major market integration effect after MERCOSUR, with a higher magnitude for Brazil-Argentina than Brazil-USA. In both cases, while PTE was greater after MERCOSUR, half-life coefficients were lower during this period. This implies that reduced arbitrage costs were accompanied by faster adjustments in price differential. It also confirms previous evidence (Barrett (2005) and Ghoshray (2010)) that a common border condition promotes faster price signals among markets. Furthermore, although PTE presented variability across products, as in the case of TCs, after the agreement this variability was reduced, suggesting that the price pairs adjust faster, regardless of the size of the shock. Overall, we managed to identify a pattern of greater price convergence for Brazil-Argentina than for Brazil-USA, confirming that there is greater market integration when trade agreements exist.

As highlighted in Liefert et al. (2010), there is an inverse relation between the domestic production and the elasticity of price signals from exporter to importer countries. This evidence is particularly representative in the case of Brazil-Argentina, in which a higher price transmission in olives (+73%) and wheat flour (+54%) was found after 1995. For the first one, since Brazil has a low domestic production, this situation could trigger a more elastic reaction of Brazilian importers during high demand periods. Similarly, for the case of wheat flour, high internal demand from the Brazilian milling industry could drive higher levels of PTE after MERCOSUR.

The highest half-life reductions after MERCOSUR on both country pairs were found on powdered milk (1.103 months) and fresh pears (2.625 months). In both cases, the domestic production of Argentina is higher than Brazil and the United States. In fact, during 2012 Argentina occupied the 4th and 6th position among the global exporters of these products respectively (FAOSTAT, 2013). It is clear that Argentina can offer these products at lower relative prices than Brazilian producers, generating and asymmetric supply behavior between both countries. These findings confirm the role played by the domestic production of the importer countries on the transmission of price signals between countries (Gilhoto and Sesso, 2010).
For the Brazil-United States pair, higher PTE changes occur for more differentiated products, such as industrial alcohol (+61%) and vegetable juices (+81%). Interestingly, in the case of wheat the PTE was higher for the United States varieties, that is, HRW1 (0.523) and SRW 15 (0.519) than for the Argentinean PAN 10 (0.441) and PAN 11 (0.422). Following Maggian and Felipe (2009), the explanations for these results could be twofold. First, the United States is a price leader for global wheat prices while Argentina is only a price follower; and second, infrastructural problems in Argentina and its lack of storage capacity, which probably limits the arbitrage activities from Argentina to Brazil.

Overall, our study highlights the importance of TCs in international trade and how it can be affected by trade agreements. The model estimates confirm that when TCs are reduced integration between markets increases, reinforcing the impact on trade. What is more, the results for the Brazil-United States pair also suggest that as long as TCs decrease, integration among markets will increase.

One explanation could be that the asymmetries among the structure of comparative and competitive advantages between both country pairs and the tariff reduction generated by MERCOSUR allows a decrease in the TCs level. Nevertheless, MERCOSUR also possesses an area related to sanitary and phytosanitary issues, that could also be relevant in the formation of total TCs. This analysis could translate into further research to determine which results could be of high relevance for policy makers.

In spite of the most relevant sources of TCs, the aforementioned results suggest that policies oriented to reducing internal trade barriers, such as having an efficient logistic infrastructure or expedite sanitary and phytosanitary inspections and procedures, could provide efficient measures to reduce TCs and impact trade in a positive manner. On the contrary, it is highly probable that having internal barriers that increase TCs will prevent full exploitation of the trade agreement, reducing its impact on member countries.

6. Concluding remarks

We used a TVECM in order analyze the effect of MERCOSUR on TCs and market integration between the Brazil/Argentina and Brazil/USA country pairs. Our findings confirmed a
significant MERCOSUR effect with lower TCs and higher PTE after the implementation of this customs area for most of the products considered in the study.

Our results suggest a positive effect on trade flows and arbitrage activities in the agri-food sectors for both country pairs, with highly heterogeneous TC and PTE variations across products. For both cases the TC reduction pattern was product-specific. It was higher for differentiated products, such as high protein wheat, powdered milk, industrial alcohol and vegetables juices, among others. With respect to PTE, the highest increases for most of the products occurred during the post-MERCOSUR period in the Brazil/Argentina pair, confirming the positive role played by distance in the transmission of price signals.

In the case of Brazil/Argentina, the lower expression of TCs could occur because of two formative components of these costs, such as variable costs and possible reductions in custom duties that effectively promoted greater efficiency in the process of Brazilian imports from Argentina. In the case of Brazil/USA, our results suggest that TC reductions occur in two forms: first, MERCOSUR could create increased investment and trade opportunities through access to a larger market, such as Brazil; and second, because MERCOSUR countries could trade products in which they do not have strong comparative advantages among themselves, and reserve trade in products in which the United States has a comparative advantage.

From our results, it was possible to conclude that MERCOSUR, despite its duty-free access, has produced fewer trade opportunities for member countries compared to non-members for some specific products. This situation is mainly because a lack of transport and communications infrastructure dampens trade opportunities and competitiveness arising from the MERCOSUR agreement. Therefore, it appears that Argentina’s membership in MERCOSUR alone is unlikely to be sufficient in overcoming its physical or economic trade barriers compared to the United States.

In conclusion, Brazil and Argentina have considerable room to maximize the benefits of MERCOSUR through the implementation of policies to develop logistics, transportation and internal distribution mechanisms. There is an opportunity to enhance competition and productivity between domestic producers and reduce the remaining barriers to external trade. Some of the more sensitive issues, such as subsidy policies, have not been addressed at a regional level, which has also affected the efficiency of the implementation of MERCOSUR. For example, the export tax in Argentina on some agricultural products has created some
distortions in trade flows with Brazil despite the existence of a duty-free area. Clearly, this diminishes the competitiveness of Argentine exports and makes room for highly competitive countries who are price makers in the global market, such as the United States. Finally, future research could focus on analyzing whether the level of subsidies affects TCs between agricultural markets, mainly focusing on the role of market structure and regulations on the integration pattern.

Acknowledgements

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Transaction cost and trade liberalization


CHAPTER 4

Driving factors of agribusiness stock markets: a panel cointegration analysis

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ABSTRACT

The recent trend towards economic regionalism has intensively promoted the establishment of trade blocs. This work explores the drivers of regional stock market integration with a focus on the agribusiness sector across the most important regional trade blocs around the world. Based on the literature on market integration and stock return pricing, we identify nine possible determinants of stock market integration, which we separate into three categories: individual market performance, macroeconomic conditions and agricultural trade. We implement panel cointegration models to analyze the stock indices of agribusiness firms in MERCOSUR, EU, APEC and NAFTA. Furthermore, we account for agriculture-specific factors to control for possible structural shifts in financial markets by including the two main commodity price bubbles during last 20 years. The results show that most of the variables included in our categories have been important factors in promoting regional stock market integration. Moreover, integration among regional stock markets was strengthened by the implementation of trade agreements. This effect is stronger in trade blocs with fewer members, such as NAFTA and MERCOSUR, compared with larger and more heterogeneous blocs such as the EU and APEC.

Keywords: stock markets, agribusiness, trade blocs, cointegration
1. Introduction

The number of Regional Trade Agreements (RTAs) increased from 34 in 1995 to 301 as of January 2013 (WTO, 2013). This situation has attracted increasing concern regarding their welfare implications and contribution to the development of stock markets (Berman et al., 2010).

While most empirical research has focused on the economic impact of trade flows between markets, recent studies have begun to address their effects on the financial sector as well. For example, Hooy et al. (2008) report that emerging stock markets in Asia have become increasingly interdependent as a result of stronger regionalism and increased liberalization. In Latin America, Carneiro and Brenes (2014) suggest that stock markets have become more regionally integrated since the implementation of trade liberalization policies in the early 1990s. These works are consistent in suggesting that stronger bilateral trade ties between countries promote a higher degree of stock market integration.

The drivers of these market integration patterns are diverse. According to Guvenen (2009), one of the main factors promoting the convergence of regional stock markets is macroeconomic integration as a result of the implementation of trade liberalization areas. He argues that greater policy coordination and market liberalization in the European Union (EU) and the Association of Southeast Asian Nations (ASEAN) has led to convergence in regional rates of return. Jawadi et al. (2010) conclude that privatization and financial deregulation policies may account for a higher level of financial integration in regional stock markets. Meanwhile, Karim and Mahid (2011) find that market-oriented policies strengthened by geographic proximity and close relationships between markets further contribute to increased stock market integration.

Initiatives to strengthen financial integration have been broadening. For example, the North American Free Trade Agreement (NAFTA), which came into force in January 1994, promoted the free transfer of all payments related to stock transactions including dividends, interest and capital gains among members (Esqueda et al., 2013). Furthermore, the ASEAN, established in August 1967, also embraced the concept of a common investment area to reduce and remove barriers to intra-regional investment among its members (Gochoco-Bautista and Remolona, 2012). In Latin America, the Common Market of the South (MERCOSUR) has implemented
transaction liberalization programs to promote quick transfers of foreign investments within their stock markets (Carneiro and Brenes, 2014).

Recent studies\(^8\) examine various effects of regional trade agreements on stock market interaction, namely: portfolio diversification, linkages in returns, volatility spillover, and more recently, contagion effects during financial crises. However, these works did not consider specific business sectors, but rather general linkages among regional stock markets.

There are very few studies on stock markets in the agricultural sector, and these have mainly focused on analyzing the effect of specific events or agricultural policies. For example, Tepe et al. (2011) investigated the impact of domestic biofuels policy on U.S. stock prices, while Pendell and Cho (2013) studied the stock market reactions by investors in Korean companies following five outbreaks of foot-and-mouth disease (FMD).

These works demonstrate regional interdependence, but do not offer insights into the driving factors of stock market integration and their implications on specific agricultural sectors. Our work aims to fill this gap by exploring the drivers of regional stock market integration with a focus on the agribusiness sector across the most important trade blocs around the world.

The objective of this paper is to explore the drivers of regional stock market integration with a focus on the agribusiness sector in several of the world’s most important international trade blocs. Based on the literature on market integration and stock return pricing, we define 9 explanatory variables, which we separate into three categories: individual market performance, macroeconomic conditions and agricultural trade. Cointegration models applied to panel data are used to analyze the stock indices of the most relevant agribusiness firms (in terms of trade) in MERCOSUR, EU, APEC and NAFTA. Episodes of high price volatility, such as price bubbles, are something specific to agricultural commodity markets which should also be accounted for when modeling the determinants of agricultural stock market integration. Accordingly, in order to account for possible structural shifts in financial markets and their impact on the behavior of agribusiness stocks, we include the two main price bubbles that have occurred during last 20 years (January 1995 to December 1996, and January 2006 to December 2010) in our analysis.

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To the best of our knowledge, none of the studies on regional stock markets have explicitly examined the impact of trade agreements together with other drivers of stock market integration on agribusiness stocks. Moreover, we also consider agriculture-specific factors to control for trade agreements effects and structural breaks, and to verify the robustness of our results using panel regression methods.

The remainder of the paper is organized as follows. Section 2 reviews the literature on the drivers of stock market integration. Section 3 and 4 describe the methodology and data respectively. Section 5 contains the empirical results and discussion. Finally, section 6 presents the conclusions of this work.

2. Drivers of regional stock market integration

Previous literature mainly focuses on modeling or analyzing the transmission process of price shocks. For example, it is widely accepted that intra-regional integration tends to be higher than inter-regional integration mainly because different time zones generate larger overlaps in trading hours between regions (Nagel and Singleton, 2011). Macroeconomic environments, contagion effects from financial crises and poor economic signals might also influence stock market integration over time. For example, Nguyen and Bhatti (2012) showed that investments yield higher conditional returns when positive indicators regarding non-farm payroll, unemployment, GDP growth and sectorial production are published.

Carrieri et al. (2007) make the first attempt to address the determinants of stock market integration. They used a pooled regression with four explanatory variables and an asset pricing approach to study the equity markets of eight emerging countries: Argentina, Brazil, Chile, India, Korea, Mexico, Taiwan and Thailand. They found that financial development and trade liberalization have a positive impact, while trade openness and global market volatility do not have a significant impact on market integration.

From an asset pricing perspective, a market integration test is commonly executed to verify the law of one price (LOP), whereby firms whose future cash flows are subject to common risks should be valued the same regardless of their location (Chami et al., 2010). Under a Capital Asset Pricing Model (CAPM) equilibrium, perfect stock market integration exists
when there are no pricing errors in benchmarking market indices with respect to a global portfolio or a list of common risk factors (Nagel and Singleton, 2011). These pricing errors could be due to limitations in common-border arbitrage, investment barriers or market inefficiency (Tepe et al., 2011).

Recent empirical work (Kose et al., 2006; Carrieri et al., 2007; Caporale and Spagnolo, 2012) suggests that the integration of regional stock markets is mainly driven by the performance of individual markets, the macroeconomic situation and trade between markets.

In particular, when a conditional asset pricing test is applied, previous authors (Bekaert et al., 2002; Dufour et al., 2010 and Nagel and Singleton, 2011) identified market development, the dividend yield differential and stock index volatility as the most relevant performance variables. A positive correlation between market development and market integration is expected because developed stock markets usually attract higher capital flows (Guangxi et al., 2012).

The dividend yield differential refers to the relationship between the domestic and global market dividend yield. It has previously been used as an instrument for evaluating the rates of return of spatially separated portfolios (Bekaert et al., 2002), and is an efficient predictor of stock integration in emerging markets because it provides clues about the relative performance of an individual market relative to global stock markets (Dufour et al., 2010).

Volatility is another important variable in explaining movements in stock returns (Guangxi et al., 2012), as studies (e.g. Grullon et al., 2012) find a positive relationship between firm-level return volatility and firm-level stock returns.

The effect of macroeconomic variables on stock market integration has been previously reported (e.g. Hilsher and Raviv, 2011). Macroeconomic stability and price indicators affect stock market integration since they influence firms’ abilities to expand their markets and, consequently, to promote investor confidence (Karim et al., 2011; Hilsher and Raviv, 2011).

The most relevant stability variables used in CAPM are exchange rate volatility and currency reserve changes (Ehrmann and Fratzscher, 2011). Exchange rate volatility is important because it affects agribusiness firms via its effect on financial returns from international trade. Changes in currency reserves have been used in international trade studies as an indicator of the economy’s ability to finance international trade (Mohanty and Turner, 2006). We assume that larger currency reserves ease firms’ financing conditions and, consequently, increase their
stock prices. Inflation and interest rates are the price indicators that most influence the stock market integration process (Aghion et al., 2009). Both affect consumption and investment costs, and as such, a firm’s expected cash flow (Ehrmann and Fratzscher, 2011). The inflation level increases financial market friction and negatively affects the efficiency of the financial system (Boyd et al., 2001). Hence, interest rates affect stock market integration by influencing capital flows between stocks and other asset markets such as bonds (Faust et al., 2007).

Agricultural goods are tradable, and many large agribusiness enterprises whose stock is publicly traded are involved in international trade. International trade, therefore, affects the cash flow of agribusiness firms, and their stock valuations. In fact, Kose et al. (2006) found a positive correlation between trade volume and stock valuation of firms participating in stock markets. In order to account for this link, we include agricultural market openness and agricultural trade intensity as explanatory variables in our model.

3. Methodology

Considering the main categories of driving forces described in the previous section, a general regional stock market integration index for a regional market \( i \) in time period \( t \) can be depicted as:

\[
R_{SI_i} = f\left( Z_{Market}, Z_{Macroeconomic}, Z_{Trade} \right)
\]

(1)

where \( R_{SI_i} \) denotes the level of regional stock integration and \( Z_{Market}, Z_{Macroeconomic}, Z_{Trade} \) are vectors of variables representing individual market performance, macroeconomic conditions and agricultural trade, respectively.

To estimate equation (1), we first require a series of estimates of \( RSI \) over time. To generate these estimates we employ a Trading-Bloc Capital Asset Price Model (TB-CAPM) (Hooy and Goh, 2008) as follows:

\[
R_{i,t} - R_{F,t} = \alpha_i + \beta_i (R_{Trade,Bloc,t} - R_{F,t}) + \varepsilon_{i,t}; \quad \forall t
\]

(2)
Driving factors of agribusiness stock markets

where $R_{i,t}$ and $R_{TradeBloc,t}$ are the returns for the market portfolio and the trade bloc portfolio respectively, and $R_{F,t}$ is the international risk free rate.

Since the pricing error $\alpha_i$ shows the deviation from perfect market integration, Levine and Zervos (1998) suggest that if stock markets are perfectly integrated, then the intercept in a regression of any asset's excess return on the appropriate benchmark portfolio (in this case $R_{TradeBloc,t} - R_{F,t}$) should be zero:

$$\alpha_1 = \alpha_2 = \ldots = \alpha_i = 0$$ (3)

Rejection of the restrictions in (3) may be interpreted as rejection of the underlying asset-pricing model or rejection of market integration. Furthermore, any significant $\alpha_i$ measures the deviation from perfect market integration and can be interpreted as the stock market integration index ($RSI_{i,t}$). $RSI_{i,t}$ can take any negative value with an upper limit of zero, where 0 indicates perfect integration.

In order to obtain a time series estimation for $\alpha_i$, we carry out a rolling regression with a set of 10 years of monthly observations. Following this estimation of $RSI_{i,t}$, we employ a panel approach with asset and time fixed effects to estimate the stock market integration model in equation (1), that we call model A:

$$RSI_{i,t} = \mu + Z_{i,t}^{T} \varphi + \eta_i + \phi_i + v_{i,t} \quad i = 1, \ldots, M; \quad t = 1, \ldots, T$$ (4)

where $\mu$ is the intercept term, $\varphi$ is a vector of $k \times 1$ coefficients and $Z_{i,t}$ is a vector of $k \times 1$ covariates for stock market $i$ and time period $t$, $\eta_i$ represents the asset fixed effects, $\phi_i$ captures the time fixed effects and $v_{i,t}$ is a random disturbance. We tested for fixed vs. random effects using the Hausman test (Hausman, 1978).

In order to test the effect of trade blocs, we estimated an additional version of the panel data model (called model B) that includes four binary variables: $DEUNI_i$, $DRAFT_i$, $DMERC_i$, and
Driving factors of agribusiness stock markets

$DAPEC_i$, representing EU, NAFTA, MERCOSUR and APEC respectively. These variables equal one if the stock market in question is located in the respective block, and zero otherwise. In order to avoid a binary variable trap, $DEUNI_i$ is chosen as the reference group. In addition, we include a second set of binary variables to represent commodity price bubbles or price turbulence in agricultural markets. The objective of this is to account for possible structural shifts in financial markets that could affect the behavior of agribusiness stocks. To define these variables, we follow Esposti and Listorti (2013) who identified the two main price bubbles that have occurred during last 20 years; the first from January 1995 to December 1996, and the second from January 2006 to December 2010. We code these as D95 and D06, respectively. Finally, all volatility series are conditional volatilities generated by estimating a GARCH process.

Accordingly, our panel model takes the following form:

$$ RSI_{it} = \mu_t + \phi_1 AMD_{it} + \phi_2 DYD_{it} + \phi_3 VOL_{it} + \phi_4 EXV_{it} + \phi_5 CRC_{it} + \phi_6 IFL_{it} + \phi_7 INT_{it} + \phi_8 AMO_{it} + \phi_9 ATI_{it} + \phi_{10} D95_{it} + \phi_{11} D06_{it} + \phi_{12} DEUNI_i + \phi_{13} D95 + \phi_{14} D06 + \phi_{15} DEUNI_i + \phi_{16} D95 + \nu_{it}; \forall i \forall t $$

(5)

4. Data description

We considered 18 stock markets in the four regional blocs with the highest volumes of agricultural trade measured in tons: the EU (European Union), MERCOSUR (Southern Common Market), NAFTA (North American Free Trade Agreement) and APEC (Asia Pacific Economic Cooperation) (Table 1). Movements in agribusiness stock prices are taken from the Morgan Stanley Capital International (MSCI) Agriculture & Food Chain Index and the GSCI Agriculture Index for each regional market from August 1993 to August 2013. To estimate the global portfolio index we use the MSCI All Country World Index as a proxy. We used the U.S Treasury bill rate obtained from the U.S. Department of the Treasury as a proxy for the world risk-free rate. The other variables in equation (5) were taken from the International Monetary Fund (IMF), Food and Agriculture Association of the United Nations (FAO), COMTRADE Database from the United Nations (UN) and the World Bank (WB). The list of our explanatory variables is presented in Table 2.
Table 1. List of trade blocs, member countries and their respective stock markets

<table>
<thead>
<tr>
<th>Country</th>
<th>EU Stock Market</th>
<th>NAFTA Stock Market</th>
<th>MERCOSUR Stock Market</th>
<th>APEC Stock Market</th>
</tr>
</thead>
<tbody>
<tr>
<td>Germany</td>
<td>Deutsche Börse</td>
<td>United States</td>
<td>Argentina</td>
<td>Chile</td>
</tr>
<tr>
<td>Spain</td>
<td>Bolsa de Madrid</td>
<td>Canada</td>
<td>Brazil</td>
<td>Australia</td>
</tr>
<tr>
<td>Italy</td>
<td>Borsa Italiana</td>
<td>Mexico</td>
<td>Uruguay</td>
<td>New Zealand</td>
</tr>
<tr>
<td>France</td>
<td>Bourse de Paris</td>
<td>Paraguay</td>
<td>BVPASA</td>
<td>Vietnam</td>
</tr>
<tr>
<td>Portugal</td>
<td>Lisbon Stock Exchange</td>
<td></td>
<td></td>
<td>Russia</td>
</tr>
</tbody>
</table>


Table 2. Description of the covariates in the stock market integration model

<table>
<thead>
<tr>
<th>Category</th>
<th>Explanatory Variable</th>
<th>Measurement</th>
<th>Reference</th>
</tr>
</thead>
<tbody>
<tr>
<td>Individual Market Performance</td>
<td>Agricultural Market Development</td>
<td>$\text{AMD} = \frac{\text{Agricultural market value}}{\text{Nominal GDP}}$</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>Dividend Yield Differential</td>
<td>$\text{DYD} = \text{DY country i} - \text{DY world}; \text{DY} = \text{dividend/price}$</td>
<td>Ang and Liu (2007)</td>
</tr>
<tr>
<td></td>
<td>Agriculture Stock Index Volatility</td>
<td>$\text{VOL} = \text{conditional volatility generated from an AR(1) process with GARCH(1,1) errors on log (Pt/Pt-1)}$</td>
<td>-</td>
</tr>
<tr>
<td>Macroeconomic Conditions</td>
<td>Exchange Rate Volatility</td>
<td>$\text{EXV} = \text{conditional volatility generated from an AR(1) process with GARCH(1,1) errors on log(ExRate). Exchange rate is expressed in terms of each domestic currency per unit of USD}$</td>
<td>Aghion et al. (2009)</td>
</tr>
<tr>
<td></td>
<td>Currency Reserve Changes</td>
<td>$\text{CRC} = \text{changes of log (international currency reserve)}$</td>
<td>Mohanty and Turner (2006)</td>
</tr>
<tr>
<td></td>
<td>Inflation Rate</td>
<td>$\text{IFL} = \frac{\text{CPI}<em>t - \text{CPI}</em>{t-1}}{\text{CPI}_{t-1}}$</td>
<td>Boyd et al. (2001)</td>
</tr>
<tr>
<td></td>
<td>Interest Reference Rate</td>
<td>$\text{INT} = \log \left( \text{Short term interest rate, TB rate or interbank rate} \right)$</td>
<td>Faust et al. (2007)</td>
</tr>
<tr>
<td>Agricultural Trade</td>
<td>Agricultural Market Openess</td>
<td>$\text{AMO} = \frac{\text{total agricultural trade with the world}}{\text{Nominal GDP}}$</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>Agricultural Trade Intensity</td>
<td>$\text{ATI} = \frac{\text{total agricultural trade with bloc members}}{\text{Total agricultural trade with the world}}$</td>
<td>-</td>
</tr>
</tbody>
</table>
5. **Empirical results and discussion**

5.1 Summary Statistics, Correlation and Panel Unit Root Tests

Table 3 presents the summary statistics for the dependent variable (RSI) and the nine covariates in the stock market integration model. The mean value of the dependent stock market integration index RSI (-0.366) and its standard deviation (0.27) suggest that significant variation in stock market integration exists among the countries that we considered.

Cointegration analysis requires that all price series are integrated to the same degree, and this can be tested by applying an appropriate unit root test. The Elliott-Rothenberg-Stock (Müller and Elliott, 2003) test results in the last column of Table 3 indicate that most of the series are I(1).

Table 3. Summary statistics and unit root test for the panel variables

<table>
<thead>
<tr>
<th>Variables</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Skewness</th>
<th>Jarque-Bera Normality Test</th>
<th>Elliot-Rothenberg-Stock (ERS) Test</th>
<th>Ho=series is non-stationarity</th>
</tr>
</thead>
<tbody>
<tr>
<td>RSI</td>
<td>-0.366</td>
<td>0.266</td>
<td>-0.000</td>
<td>-1.937</td>
<td>-1.366</td>
<td>3124.1**</td>
<td>-2.542**</td>
<td></td>
</tr>
<tr>
<td>AMD</td>
<td>0.011</td>
<td>0.254</td>
<td>3.496</td>
<td>-2.674</td>
<td>-0.290</td>
<td>88675.3*</td>
<td>-1.642</td>
<td></td>
</tr>
<tr>
<td>DYD</td>
<td>0.0039</td>
<td>0.008</td>
<td>0.040</td>
<td>-0.018</td>
<td>0.628</td>
<td>444.7*</td>
<td>-1.338</td>
<td></td>
</tr>
<tr>
<td>VOL</td>
<td>8.804</td>
<td>1.019</td>
<td>13.831</td>
<td>0.811</td>
<td>4.093</td>
<td>177653.1*</td>
<td>-1.115</td>
<td></td>
</tr>
<tr>
<td>EXV</td>
<td>0.001</td>
<td>0.021</td>
<td>1.229</td>
<td>0.001</td>
<td>41.254</td>
<td>23488.7**</td>
<td>-1.813</td>
<td></td>
</tr>
<tr>
<td>CRC</td>
<td>0.001</td>
<td>0.064</td>
<td>0.671</td>
<td>-0.669</td>
<td>-0.908</td>
<td>77645.3**</td>
<td>-2.958**</td>
<td></td>
</tr>
<tr>
<td>IFL</td>
<td>0.002</td>
<td>0.105</td>
<td>2.823</td>
<td>-4.200</td>
<td>-6.543</td>
<td>9991.7*</td>
<td>-0.669</td>
<td></td>
</tr>
<tr>
<td>INT</td>
<td>-2.547</td>
<td>0.649</td>
<td>-0.091</td>
<td>-5.872</td>
<td>-0.317</td>
<td>656.1**</td>
<td>-1.665</td>
<td></td>
</tr>
<tr>
<td>AMO</td>
<td>0.223</td>
<td>0.365</td>
<td>2.054</td>
<td>0.000</td>
<td>2.271</td>
<td>17984.8*</td>
<td>-1.650</td>
<td></td>
</tr>
<tr>
<td>ATI</td>
<td>0.387</td>
<td>0.220</td>
<td>0.712</td>
<td>0.000</td>
<td>-0.371</td>
<td>333.2*</td>
<td>-1.471</td>
<td></td>
</tr>
</tbody>
</table>

Note 1: The asymptotic critical values are -3.48***, -2.89** and -2.57* for the ERS Test
Note 2: *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively
The matrix of correlation coefficients between covariates (Table 4) indicates that there is little multicollinearity among them. With respect to the correlation index, only 9 coefficients exceed the value 0.1, with 0.31 being the highest coefficient (EXV/VOL relation).

Table 4. Correlation matrix for Panel Series of the Model

<table>
<thead>
<tr>
<th></th>
<th>RSI</th>
<th>AMD</th>
<th>DYD</th>
<th>VOL</th>
<th>EXV</th>
<th>CRC</th>
<th>IFL</th>
<th>INT</th>
<th>AMO</th>
<th>ATI</th>
</tr>
</thead>
<tbody>
<tr>
<td>RSI</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>AMD</td>
<td>0.0005*</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>DYD</td>
<td>0.0451</td>
<td>-0.0227</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>VOL</td>
<td>-0.2191*</td>
<td>-0.0245</td>
<td>-0.0381</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>EXV</td>
<td>-0.0513*</td>
<td>-0.0009</td>
<td>0.0016</td>
<td>0.3140**</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>CRC</td>
<td>-0.0046</td>
<td>0.0613*</td>
<td>-0.0234</td>
<td>0.0280</td>
<td>-0.0117</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>IFL</td>
<td>-0.0091</td>
<td>-0.0005</td>
<td>-0.0032</td>
<td>0.0201</td>
<td>0.0179</td>
<td>-0.0124</td>
<td>1.0000</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>INT</td>
<td>0.0957*</td>
<td>-0.0368</td>
<td>0.0360</td>
<td>0.3097**</td>
<td>0.1035*</td>
<td>0.0544*</td>
<td>0.0158</td>
<td>1.0000</td>
<td></td>
<td></td>
</tr>
<tr>
<td>AMO</td>
<td>0.0026</td>
<td>0.0432</td>
<td>-0.0116</td>
<td>-0.1161*</td>
<td>-0.0222</td>
<td>0.0045</td>
<td>-0.0047</td>
<td>-0.2719**</td>
<td>1.0000</td>
<td></td>
</tr>
<tr>
<td>ATI</td>
<td>0.0910</td>
<td>-0.0127</td>
<td>0.0086</td>
<td>-0.1737*</td>
<td>-0.0432</td>
<td>-0.0374</td>
<td>0.0061</td>
<td>0.0333</td>
<td>-0.0440</td>
<td>1.0000</td>
</tr>
</tbody>
</table>

Note: *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively.

5.2 Estimated Results for the Panel Model

We conducted a series of Hausman and F tests to determine whether random or fixed effects were present. Among all specifications, we found that the two-way fixed effect model presents the highest adjusted $R^2$ and lowest residual sum of squares (RSS) (Table 5).
Table 5. Hausman and F tests for model selection of the panel regression

<table>
<thead>
<tr>
<th>Hypothesis</th>
<th>Adj R²</th>
<th>RSS</th>
<th>Chi-Sq</th>
<th>F</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Hausman Test for Random Effects</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( H_0 ): One-Way Cross-section Random Effects</td>
<td>0.03</td>
<td>255.96</td>
<td>15.98</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): One-Way Cross-section Fixed Effects</td>
<td>0.29</td>
<td>246.69</td>
<td>0.119</td>
<td></td>
</tr>
<tr>
<td>( H_0 ): One-Way Period Random Effects</td>
<td>0.08</td>
<td>327.24</td>
<td>71.15</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): One-Way Period Fixed Effects</td>
<td>0.11</td>
<td>310.98</td>
<td>0.013**</td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Two-Way Random Effects</td>
<td>0.04</td>
<td>235.46</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Cross-section Random Period Fixed Effects</td>
<td>0.12</td>
<td>221.42</td>
<td>0.101*</td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Two-Way Random Effects</td>
<td>0.04</td>
<td>235.46</td>
<td>Failed</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Cross-section Fixed Period Random Effects</td>
<td>0.29</td>
<td>233.78</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Two-Way Random Effects</td>
<td>0.04</td>
<td>235.46</td>
<td>Failed</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.34</td>
<td>219.24</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Two-Way Cross-section Random Period Fixed Effects</td>
<td>0.12</td>
<td>221.42</td>
<td>31.00</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.33</td>
<td>219.24</td>
<td>0.012**</td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Two-Way Cross-section Fixed Period Random Effects</td>
<td>0.29</td>
<td>233.78</td>
<td>144.10</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.33</td>
<td>219.24</td>
<td>0.012**</td>
<td></td>
</tr>
<tr>
<td><strong>F-Test for fixed effects</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( H_0 ): Without fixed effects</td>
<td>0.22</td>
<td>402.99</td>
<td>55.15</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): One-Way Cross-section Fixed Effects</td>
<td>0.44</td>
<td>312.13</td>
<td>0.037**</td>
<td>0.050**</td>
</tr>
<tr>
<td>( H_0 ): Without Fixed Effects</td>
<td>0.22</td>
<td>501.34</td>
<td>2.45</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): One-Way Period Fixed Effects</td>
<td>0.16</td>
<td>399.89</td>
<td>0.013**</td>
<td>0.000***</td>
</tr>
<tr>
<td>( H_0 ): Without Fixed Effects</td>
<td>0.20</td>
<td>415.87</td>
<td>13.34</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.67</td>
<td>323.98</td>
<td>0.001**</td>
<td>0.000***</td>
</tr>
<tr>
<td>( H_0 ): One-Way Cross-section Fixed Effects</td>
<td>0.43</td>
<td>303.19</td>
<td>83.99</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.56</td>
<td>299.00</td>
<td>0.000***</td>
<td>0.000***</td>
</tr>
<tr>
<td>( H_0 ): One-Way Period Fixed Effects</td>
<td>0.16</td>
<td>415.34</td>
<td>4.65</td>
<td></td>
</tr>
<tr>
<td>( H_1 ): Two-Way Fixed Effects</td>
<td>0.51</td>
<td>304.00</td>
<td>0.000***</td>
<td>0.000***</td>
</tr>
</tbody>
</table>

Note 1: *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively.
Note 2: * Test failed as the test variance (either cross-sectional or period) is invalid.
As previously explained, we estimated an additional version of the panel data model (called model B) in which specific binary variables are introduced in the panel regression. We replaced the cross-section and period fixed effects with the trade agreement and price bubble binary variables respectively. In this case, the cross-section fixed effects are replaced by the trade bloc binary variables described above. Furthermore, to test whether structural shifts in financial markets due to episodes of price bubbles have influenced agribusiness stock market integration, we replace the period fixed effects in the second model with the D95 and D06 binary variables. The estimation results for both models are presented in Table 6.

Table 6. Estimation results for the models A and B

<table>
<thead>
<tr>
<th></th>
<th>Model A coefficients</th>
<th>Model B coefficients</th>
</tr>
</thead>
<tbody>
<tr>
<td>DEU</td>
<td>-0.3876**</td>
<td>-0.5773**</td>
</tr>
<tr>
<td>AMD</td>
<td>0.0012**</td>
<td>0.0029</td>
</tr>
<tr>
<td>DYD</td>
<td>0.6155</td>
<td>-0.2865</td>
</tr>
<tr>
<td>VOL</td>
<td>-0.0097***</td>
<td>-0.0123***</td>
</tr>
<tr>
<td>EXV</td>
<td>-0.0088</td>
<td>-0.3255**</td>
</tr>
<tr>
<td>CRC</td>
<td>-0.0875</td>
<td>-0.0451</td>
</tr>
<tr>
<td>IFL</td>
<td>0.0134</td>
<td>-0.0118*</td>
</tr>
<tr>
<td>INT</td>
<td>0.0883</td>
<td>0.2421**</td>
</tr>
<tr>
<td>AMO</td>
<td>-0.1977**</td>
<td>-0.0128***</td>
</tr>
<tr>
<td>ATI</td>
<td>0.3888**</td>
<td>0.4561**</td>
</tr>
<tr>
<td>D95</td>
<td></td>
<td>0.0499***</td>
</tr>
<tr>
<td>D06</td>
<td></td>
<td>-0.0491***</td>
</tr>
<tr>
<td>DEUNI</td>
<td>0.3162**</td>
<td></td>
</tr>
<tr>
<td>DNAFT</td>
<td>0.3703**</td>
<td></td>
</tr>
<tr>
<td>DMERC</td>
<td>0.3523***</td>
<td></td>
</tr>
<tr>
<td>DAPEC</td>
<td></td>
<td>0.2967***</td>
</tr>
<tr>
<td>Adj R2</td>
<td>0.4521</td>
<td>0.4188</td>
</tr>
<tr>
<td>RSS</td>
<td>291.2434</td>
<td>398.3212</td>
</tr>
</tbody>
</table>

Note: *, ** and *** denote significance at the 0.10, 0.05 and 0.01 levels, respectively

In Model A four variables were found to be statistically significant, compared with 12 in Model B. The magnitudes of the estimated coefficients are highly consistent in both models, except for the dividend yield differential, inflation rate and agricultural market openness.
5.2.1 Individual market performance category

In both models we found a positive relationship between RSI and the size of the agribusiness sector as a percentage of total GDP (AMD). This result suggests that increased capitalization of agribusiness firms as a percentage of GDP stabilizes the firm’s valuation. The previous statement confirms the notion that high sector growth expectations would promote market integration via greater financial stability. In emerging markets, the relationship between sector GDP growth and stock valuation has been demonstrated previously by the work of Pendell and Cho (2013).

The market volatility coefficients (VOL) are \((-0.0097\) for model A and \((-0.0123\) for model B), suggesting that this has a negative effect on the integration process between stock markets. Increased levels of volatility do not promote regional stock market integration because this mainly depend of the development level of each stock market, and is driven by a different economic process than pricing returns (Bekaert and Harvey, 1997). In fact, Esqueda et al. (2013) demonstrated that the correlation between local and global market returns increases as a result of market-oriented policies, but these policies do not drive up local market volatility. In both models, DVD shows statistically significant coefficients, but in different directions (positive in model A and negative in model B). This impedes us from being able to make conclusions regarding a single effect of this variable on RSI.

Certainly, the aforementioned results confirm that countries can benefit from trade regionalism policies in terms of stock return pricing, by acting as an economic bloc rather than as individual markets. The main consequence of such policies is an increase in market options and stability as a result of higher capitalization levels.

5.2.2 Macroeconomic conditions category

This category includes macroeconomic and stability indicators, namely: exchange rate volatility (EXV), inflation rate (IFL), interest reference rate (INT) and currency reserves changes (CRC).

As expected, EXV has a significant and negative effect on agribusiness stock market integration in both models \((-0.088\) and \(-0.3255\) for model A and B respectively), confirming an inverse relationship with respect to stock market integration. According to Gounopoulos et al. (2013), unexpected changes in currency values affects price convergence among stock markets because of their effect on macroeconomic stability and economic expectations. Both coefficients are high
Driving factors of agribusiness stock markets

in absolute value, suggesting that this variable has an important impact as driving factor of regional stock markets. Because it affects the economy’s ability to finance international trade, the large-scale use of currency reserves has significant macroeconomic implications. In fact, our model’s coefficients show an inverse relationship between changes in currency reserves (CRC) and stock market integration. Changes in currency reserves are often associated with exchange rate changes that increase the risks of holding instruments. This has a positive relationship with the risk exposures of stock instruments due to the use of foreign exchange reserves to resist currency appreciation, which has been demonstrated previously by Mohanty and Turner (2006). INF has a negative coefficient in both models, but only in model B is it statistically significant. Bittencurt (2011) finds an inverse relationship between inflation and stock market integration because high inflation levels intensify financial market frictions and reduce the efficiency of the financial system. Moreover, our results are in line with those of De Grauwe (2012) who demonstrated that high inflation rates dampen the attraction of capital flows, further decreasing arbitrage activities between stock markets. Interest rates represent the return on alternative assets to stocks, and the discount rates used on stock return valuations. In models A and B, INT is positively correlated with market integration, recording values of 0.09 and 0.24 respectively. Ehrmann and Fratzscher (2011) conclude that higher interest rates increase stock market integration because they divert capital from the stock market to the bond market. With the increasing ease of market access as a result of trade agreements, regional stock investments are promoted because higher bond returns attract local investors to introduce their capital into regional markets.

5.2.3 Agricultural trade category

The trade indicators used were agricultural market openness (AMO) and agricultural trade intensity (ATI). As expected, agricultural trade intensity (ATI), which measures total agricultural trade with block members as a proportion of total agricultural trade with the world, unlike agricultural market openness (AMO) which measures agricultural trade with the world as a proportion of nominal GDP, is positive and significant in both models. While the ATI values were 0.3888 for model A and 0.4561 for model B, the AMO coefficient was -0.1977 and -0.01288 respectively.
Driving factors of agribusiness stock markets

ATI has a positive relationship with RSI because the higher the trade flow between regional bloc members, the higher the level of regional stock market integration. This implies that higher levels of market access and trade due to regional liberalization policies promote stock market integration and thus capital flows between countries or sectors, further supporting the hypothesis of this work.

Although previous studies on developed countries (Aghion et al., 2009 and De Grauwe, 2012) suggested that volatility and/or risk are important factors in the behavior of stock markets, our results demonstrate that for the agribusiness sector the effect of this type of determinant variable is not so significant. This could benefit emerging countries or developing markets that usually present higher risk qualifications. Overall, we confirm the relevance of the agribusiness sector within the country, and economic blocs are key driving factors of stock market integration. These results constitute a novel contribution to the literature, as most studies have not considered sectorial characteristics, production and trade among the factors that promote the integration of stock markets.

5.2.4 Price bubbles and Trade blocs

Price bubbles are driving factors of market integration because of their effects on price differences and consequently arbitrage activities (Gilbert, 2010). In our model, all price bubble variables present statistically significant coefficients. This result supports the existence of two-way fixed effects. Among these, the first period dummy (D95) had a positive and significant effect on market integration with an estimated coefficient of 0.0499. However, the second period (D06) had a negative effect of -0.0491. The higher impact on market integration from the D95 bubble could be a consequence of the increased arbitrage activities among commodity traders when compared to 2006 bubble because of the higher domestic/international price spread during 1995 bubble when compared to 2006 (Gilbert, 2010; Esposti and Listorti, 2013). The D95 bubble had the effect of increasing stock prices and as such return pricing behavior for agribusiness stocks. In other words, the greater price spread during the 1995 crisis seems to have promoted arbitrage activities among regional investors who held agribusiness stocks from markets with lower commodity prices to markets with higher commodity prices.

The last group of variables included in model B is the presence of regional trade agreements pertaining to each of the countries considered in this work. For both models, they present positive
and statistically significant coefficients. Our findings provide a clear picture of the positive effect of the implementation of trade agreements on regional market integration, further confirming previous results (Gochoco-Bautista and Remolona, 2012; Caporale and Spagnolo, 2012). The impact on market integration is relatively homogeneous for all agreements, suggesting that regardless of the level the economic development of each country and stock market price arbitrage among different markets, the benefit in all blocs is present to a similar degree.

The highest degree of integration is found in a developed market, NAFTA, followed by MERCOSUR, EU and APEC. In the case of NAFTA and MERCOSUR, factors such as distance and public policies could drive a greater degree of market integration. This argument is supported by Lahrech and Sylwester (2013), who find a positive relationship between the integration of financial markets and the distance between them. However, our results for the EU - where the markets considered in Germany, Italy, France, Portugal and Spain are geographically closer to one another than the markets in, for example NAFTA - are more similar to those of Bekaert et al. (2011). Their results demonstrate an inverse relationship between the number of stock markets and the degree of pricing efficiency when a large number of stock markets are analyzed. Bekaert et al. (2011) suggest that financial provisions policies to reduce transaction costs for capital flows can promote greater integration in regional risk pricing. Recently, both blocks have generated legal initiatives with the aim of fulfilling this objective and promoting capital flows between their member countries.

Furthermore, the heterogeneity of the members of an RTA might play a more important role than geographic distance between them. For example, in the case of APEC the differences between their stock markets in terms of trade volume, the distance (for example, between Chile and Russia) and the competitive trade policies among them could work against effective cooperation between countries. The negative effect of these policies on capital flows and financial cooperation has been demonstrated previously in the works of Click and Plummer (2005) and Karim and Ning (2013).

Overall, our findings suggest that trade blocs as well as price bubbles play a relevant role in arbitrage activities between stock markets. Moreover, market stability indicators (i.e. AMD and/or VOL) have more explanatory power than performance level, such as for example DYD, as drivers of regional stock market integration. The integration of stock markets is more exposed to individual and regional market stability than world level conditions. Finally, the empirical
evidence documented here demonstrates that market stability is a significant factor for convergence in regional risk pricing on stock markets.

6. Conclusion

The recent trend towards economic regionalism and its effect on the agricultural sector has become a catalyst for research on the stock dynamics of agribusiness firms. This work aimed to explore the drivers of regional stock market integration with a focus on the agribusiness sector across the most important trade blocs around the world. This was a previously unexplored issue in the literature on stock market integration.

In this study, a regional integration framework, which combines information from three categories and nine explanatory variables, is used to examine the contributing factors to regional stock market integration. We also consider agriculture-specific factors to control for possible structural breaks and trade agreement effects.

We conclude that our categories have been important factors in promoting regional stock market integration, further suggesting that this process is driven by the market performance, macroeconomic conditions and agricultural trade of each country. Among them, market stability indicators and intra-regional trade levels have greater explanatory power than performance variables as drivers of regional stock market integration. Besides this, the integration of stock markets is more exposed to individual and regional market stability than world level conditions (for example ATI vs AMO). Furthermore, stock market integration was strengthened by the implementation of trade agreements and during the 1995 price bubble.

The level of market integration differs according the number of countries participating in the each stock market. The integration level is higher for small trade blocs (in terms of countries), such as for example NAFTA and MERCOSUR, when compared with a more segmented bloc such as the EU or APEC. Our results suggest that the level of development and public policies also play a role in the integration pattern of each regional market.

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Driving factors of agribusiness stock markets


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CHAPTER 5.

Spatial market integration and fuel prices: an empirical analysis of the Chilean horticultural sector

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ABSTRACT

The role played by the fuel prices on the price transmission between central and regional horticultural markets has been an unexplored field of research on developing countries. The aim of this article is to propose a procedure to estimate a regime-dependent vector error correction model with an exogenous threshold variable (Chilean retail fuel prices) were not only the short and long-run equilibrium relationship itself can display threshold-type non-linearity. The proposed approach is unique in explicitly testing the threshold cointegration process based on the Gonzalo and Pitarakis (2006) test. We considered the most relevant central and regional wholesale markets (Santiago and Talca) the prices of the most planted Chilean horticultural products, namely: maize, tomato, onion, carrot. In order to account for the fuel prices, we consider the Chilean retail prices. The research was conducted using a price series of weekly frequency for the period January 2009 to December 2013. We identify two price transmission regimes characterized by different equilibrium relationships and short-run adjustment processes. For mostly products, our empirical results confirms an unidirectional price transmission process from central to regional markets which confirms a symmetric price adjustment and emphasized the role of fuel prices on the price and speed of adjustments transmission elasticity. The results imply that formation of the wholesale prices in both of the analyzed markets tend to be product specific. The method proposed by this work appers to be is particularly suitable for capturing irregular seasonal threshold effects in price transmission which is typical for vegetables.

Keywords: Spatial price transmission, market integration, threshold cointegration
1. Introduction

Price is the primary mechanism by which various levels of the market are linked. The extent of adjustment and speed with which shocks are transmitted among producer, wholesale, and retail market prices is an important factor reflecting the actions of market participants at different levels. In this sense, the transmission of changes in the producer price to changes in the consumer price depends, however, greatly on the type of product. For example, products that are perishable and undergo minimal processing such as vegetables, fruit, and fresh milk, are expected to have a relatively quick price transmission mechanism. Markets that are not integrated may convey inaccurate price information, leading to misguided policy decisions and a misallocation of resources. Sexton et al., (1991) identified three reasons for a lack of market integration: imperfect competition, different trade barriers and prohibitive transactions costs. Transaction costs are composed of transport cost, variable costs, customs duties and unmeasurable costs. Among them, the most relevant in terms of price formation for all traded agricultural products are the transport costs that are highly dependent of fuel prices (Ejrnæs and Persson, 2000). Transport costs loom large in Chilean markets, because the long distances between main wholesale markets and transport infrastructure dependent primarily on trucks. The movement of products between warehouse markets provides a straightforward observation of large variations in transport distance, given that many fruits and vegetables are produced domestically. Horticultural products are particularly sensitive to changes in fuel price because of the need for refrigeration during transport. As a result, given that these products undergoes minimal processing and packaging, the most important component of transaction cost is likely to be fuel used in transportation. The effect of fuel prices on transport cost is well-identified, fuel price increases can drive up transport costs, which in turn affect the price of agricultural products (Eryigit and Karaman, 2011). Accordingly, the role played by fuel prices offers an appropriate method to analyse the effect of transport cost on the spatial integration level between horticultural wholesale markets. In fact, there is still no clear evidence on the effect of the fuel prices has on the integration process of horticultural markets in Chile. Consequently, the objective of this paper is to analyze the role played by this factor on the spatial market integration between the major horticultural products historically traded between two representatives wholesale markets Santiago (a central market) and Talca (a representative regional market).
The use of regime dependent models to capture transaction costs are diverse, for example, Negassa and Myers (2007) use parity bounds model to analyse policy effects and spatial market efficiency using threshold model. Bakucs et al. (2007) and Bruemmer et al. (2009) apply Markov Switching Models to analyse horizontal integration of German and Hungarian wheat prices, and wheat prices in Ukraine respectively. Moreover, Bakucs and Ferto (2008) use Hansen and Seo (2002) threshold cointegration test and Threshold Vector Error Correction model to test regional milk price integration in Hungary.

Applications of regime dependent models assume that prices are linked by a constant long-run equilibrium relationship (see for example, Alam et al. 2012). In the horticultural sector this assumption may not always be justifiable. Horticultural markets are characterized by a fixed-proportion relationship among farm inputs, wholesale products, and final retail products, as well as an inelastic short-run supply owing to perishability and the timing of growing seasons (Pingali et. al 2005). In order to overcome this issue, in this paper we propose a three-step procedure to estimate a regime-dependent vector error correction model (VECM). In this model, not only the short-run adjustment process towards equilibrium, but also the long-run equilibrium relationship itself can display threshold-type non-linearity, as a function of the size of a stationary variable, that in our case is the fuel prices, with respect to a threshold value. The proposed approach is unique in explicitly testing the null hypothesis of linear cointegration against the alternative of threshold cointegration based on a test proposed by Gonzalo and Pitarakis (2006). We hypothesized, that long-run price relationships between Santiago and Talca wholesale horticultural markets may display different patterns depending on the fuel prices.

The article proceeds as follows: the next section review relevant literature on market integration. Sections 3 and 4 describe the corresponding econometric methodology and data sources. Section 5 and 6 present the results and discussion, respectively. Lastly, section 7 gives a conclusion.

2. The spatial market integration

The magnitude of trade relationship between markets or regions implies the existence of some kind of price coordination mechanism (Bakucs and Ferto 2014). Perfectly integrated markets are usually assumed to be efficient as well. In this paper however, we analyze whether long-run price relationships underlying trade relationships exist between two horticultural wholesale markets
Tomek and Robinson (2003), defines the two axioms of the regional price differences theory:

1. The price difference in any two regions or markets involved in trade with each other equals the transaction costs.
2. The price difference between any two regions or markets not involved in trade with each other is smaller than the transaction costs.

Let’s consider, two spatially different markets, where the price of a given good in time $t$ is $P_{1t}$ and $P_{2t}$ respectively. The two markets are considered integrated, if the price on market 1 equals the price on market 2 corrected with transport costs, $K_t$:

$$P_{1t} = P_{2t} + K_t$$  \hspace{1cm} (1)

Accordingly, trade between the two markets occurs only if $|P_{1t} - P_{2t}| > K_t$. That means that the arbitrage ensures that prices of the same good traded in spatially separate markets equalise or converge.

Following the definition given by Barrett (2001), TCs between a market “i” and “j” are composed of transport costs ($f_{ij}$), where distance is one of the most important factors in the how TCs are formed; variable costs ($v_{ij}$) associated with rates, cargo insurance, contracts, financial expenses, hedging, sanitary and phytosanitary barriers; customs duties ($d_{ij}$); and unmeasurable costs ($w_{ij}$), such as opportunity cost, the cost of searching for information and risk premiums.

Fuel prices can affect agricultural sector through different channels, being the most important their link through transport costs. Fuel price increases can drive up transport costs, which in turn affect the prices of all traded agricultural products. Interest in how fuel prices may affect fruits and vegetable prices spans several areas of policy-oriented economic research. For example, questions abound as to why retail food prices have become considerably more volatile since 2006 (Tegene, 2009) and if fuel price rallies affects the welfare level of rural producers due their effect on transaction cost (Dillon and Barrett, 2013)

Transport costs loom large in Chilean markets, because the long distances between main wholesale markets and transport infrastructure dependent primarily on trucks. The movement of produce between warehouse markets provides a straightforward observation of large variations in
transport distance, given that many fruits and vegetables are both produced domestically and imported. Fresh produce is particularly sensitive to changes in fuel price because of the need for refrigeration during transport (Pede, 2005). As a result, given that fresh produce undergoes minimal processing and packaging, the most important component of transaction cost is likely to be fuel used in transportation.

Applications of regime dependent models of price transmission assume that prices are linked by a constant long-run equilibrium relationship, while allowing for threshold or switching effects in the short-run adjustment process towards this equilibrium (Gonzalo and Pitarakis, 2006). In the horticultural sector this assumption may not always be justifiable. Horticultural markets are characterized by a fixed-proportion relationship among farm inputs, wholesale products, and final retail products, as well as an inelastic short-run supply owing to perishability and the timing of growing seasons (Goetz and Von Cramon-Taubadel, 2008). In this context, the factors that might contribute to dampen the transmission of price signals are: inadequate infrastructure, transportation bottlenecks, lack of market information, information asymmetry, market power, menu cost (Park et al. 2002). These kinds of factors are common in developing countries agricultural markets such as Chile and pose serious challenges to policy makers. So, estimating the threshold in the price adjustment from one market to another or from one level to another in the supply chain should be a rule rather than an exception, especially in the context of horticultural markets.

3. Methods

The method proposed by Gonzalo and Pitarakis (2006) will be used to test the linear cointegration null against the threshold cointegration between markets of Santiago and Talca. As required, we will use the Chilean fuel price series as exogenous threshold variable. In this approach, the null hypothesis of linear cointegration are:

\[ y_t = \beta' x_t + u_t, \quad (2) \]

against the alternative hypothesis of cointegration with threshold effects:

\[ y_t = \beta' x_t + \lambda' x_t I(q_{t-d} > \gamma) + u_t, \quad \text{with} \ x_t = x_{t-1} + v_t, \quad (3) \]
where $u_t$ and $v_t$ are scalar and p-vector valued stationary disturbance terms respectively, $q_{t-d}$ with $d \geq 1$ is a stationary threshold variable lagged by $d$ periods, and $I(q_{t-d} > \gamma)$ is an indicator function that equals one if $q_{t-d} > \gamma$, and zero otherwise.

Gonzalo and Pitarakis (2006) proposed a supLM test based on the following statistic:

$$LM_T(\gamma) = \frac{1}{\sigma^2_0} u'MX(\gamma)(X'MX(\gamma))^{-1}X'Mu$$

(4)

where $M = I - X(X'X)^{-1}X'$, $X$ stacks all values of $x_t$ in the linear model (1), and $X(\gamma)$ stacks the values of $x_t$ corresponding to the criterion $q_t > \gamma$ in the non-linear model (2). $T$ is the length of the full sample, $u$ is the residual, and $\sigma^2_0$ is the residual variance of the linear model (1).

The LM test statistic $LM_T(\gamma)$ is calculated for all possible values of the threshold variable $q_t$. A trimming parameter is employed to ensure a minimum number of observations on each side of the threshold. The supLM test statistic is given by:

$$supLM = sup_{\gamma \in T} LM_T(\lambda)$$

(5)

Critical values for this test statistic are taken from Andrews (1993).

The next step is to estimate an unrestricted, regime-specific ECM by including dummy variables defined by the indicator function $I(q_{t-d} > \gamma)$ corresponding to the threshold determined by the supLM test. This ECM takes the form:

$$\Delta y_t = \beta_0 + \delta_0 * I(q_{t-d} > \gamma) + \sum_{n=1}^{K} (\beta_{2n} \Delta x_{t-n+1} + \delta_{1n} \Delta x_{t-n-1} \times I(q_{t-d} > \gamma)) - \sum_{n=1}^{K} (\beta_{2n} \Delta y_{t-n} + \delta_{2n} * I(q_{t-d} > \gamma)) + \beta_1 * y_{t-1} + \delta_1 * y_{t-1} \times I(q_{t-d} > \gamma) + \beta_4 * x_{t-1} + \delta_4 * x_{t-1} \times I(q_{t-d} > \gamma) - \epsilon_t$$

(6)

The regime-dependent cointegration vector can be retrieved from equation (5) as:

$$\alpha_0 = - (\beta_0 + \delta_0 * I(q_{t-d} > \gamma)) / (\beta_3 + \delta_3 * I(q_{t-d} > \gamma)) \text{ and}$$

$$\alpha_1 = - (\beta_4 + \delta_4 * I(q_{t-d} > \gamma)) / (\beta_3 + \delta_3 * I(q_{t-d} > \gamma))$$

(7)

(8)
Accounting for this type of non-linearity in a cointegration regression allows the long-run relationship to move back and forth between regimes as a function of a threshold variable, rather than hypothesising a one-off break in this relationship. This is an appealing model in settings, such as the one explored below (spatial trade in the horticultural sector), in which price transmission is hypothesised to be seasonal, but the timing and duration of seasons differs from year to year depending on weather and harvests in the regions that are linked by trade. In such settings, the use of seasonal dummy variables to account for seasonal variation in the equilibrium relationship (e.g. Chavas and Mehta, 2004) might not be sufficiently flexible.

4. Data Description

The research was conducted with a price series of weekly frequency for the period January 2009 to December 2013. We considered the wholesale prices (in US Dollars/Kg) for the most relevant Chilean horticultural products (in area planted), namely: maize, tomato, onion and carrot. In Chile, these products account for 38% of the total surface, 28% of the total domestic trade (in value) and 18% of the total producer units (ODEPA, 2014). In order to account for the fuel prices, we consider the Chilean retail prices which are reported in Chilean Pesos (CHL), however, in order to facilitate the analysis, and to avoid inflation induced trends, data was transformed into Dollars per liter using monthly average US Dollar per CHL, the exchange rate provided by the Chilean Central Bank. Table 1 offers a description of the variables.

Table 1. Description of the variables considered in this study

<table>
<thead>
<tr>
<th>Variable</th>
<th>Market (City)</th>
<th>Planted Area (Hectares)</th>
<th>Participation on Total Surface (%)</th>
<th>Variable description</th>
</tr>
</thead>
<tbody>
<tr>
<td>Maize</td>
<td>Santiago, Talca</td>
<td>13558</td>
<td>17</td>
<td>Wholesale price</td>
</tr>
<tr>
<td>Tomato</td>
<td>Santiago, Talca</td>
<td>7253</td>
<td>9</td>
<td>Wholesale price</td>
</tr>
<tr>
<td>Onion</td>
<td>Santiago, Talca</td>
<td>6742</td>
<td>7</td>
<td>Wholesale price</td>
</tr>
<tr>
<td>Carrot</td>
<td>Santiago, Talca</td>
<td>5463</td>
<td>5</td>
<td>Wholesale price</td>
</tr>
<tr>
<td>Fuel Sale Price</td>
<td>Chilean retail price</td>
<td></td>
<td></td>
<td>Exogenous variable</td>
</tr>
</tbody>
</table>
5. Results

The empirical analysis is done in three steps. First, we assess the time series properties of data by carry out a unit root test by applying an ADF test (Dickey-Fuller, 1979) and KPSS test (Kwiatowski et al., 1992). They suggest that the prices in Santiago ($p_S$) and Talca ($p_T$) are I(1)\(^9\). Second, we test the cointegration degree of the series by using a Johansen test (Johansen, 1988). This suggest that there is cointegration over the whole sample, thus allowing to implement the Gonzalo and Pitarakis (2006) cointegration test. Accordingly, we proceed to test the importance of trade flows between the horticultural markets of Santiago and Talca. More exactly, whether the fuel prices plays any role in horizontal spatial price market integration between a Central and a Regional horticultural wholesale market. We conduct the Gonzalo and Pitarakis (2006) procedure for threshold cointegration between $p_S$ and $p_T$ for model (I) $p_T = f(p_S)$ with $p_T$ as the dependent and $p_S$ as the independent variable. The Chilean retail fuel prices is used as the threshold variable and since this variable fluctuates from day to day, we smooth it by calculating the central moving average of the nearest 24 observations for each observation. For all variables, the LM-test statistic in (5) is estimated for all observed values of the threshold variable, with the trimming parameter is set to 0.05 to ensure that each regime contains at least 5\% of all observations.

\(^9\) All unit root and cointegration tests are available upon request
Figure 1. Values of the Gonzalo-Pitarakis sup LM$\gamma$ Statistics with respect to the Chilean Fuel Prices.

a) Maize

b) Tomato
According to our estimation, for all variables two regimes with distinct long-run parameters were identified (see figure 1: a-d).

For maize, the maximum $LM_{\gamma}$ statistics corresponds to a threshold value of 17.834. The first regime contains 79 observations and is active when the price of fuel is less than 561 Chilean pesos/liter (0.91 USD). The second, larger regime pools 98 observations, and is active when the fuel price is more than 561 Chilean pesos/liter (0.91 USD).
For tomato, the maximum $LM_{\gamma}$ statistics corresponds to a threshold value of 17.877. The first regime contains 101 observations, and is active when the price of fuel is less than 625 Chilean pesos/liter (1.02 USD). The second, larger regime pools 73 observations, when the fuel price is more than 625 Chilean pesos/liter (1.02 USD).

For carrot, the maximum $LM_{\gamma}$ statistics corresponds to a threshold value of 6.858. The first regime contains 117 observations, and is active when the price of fuel is less than 661.9 Chilean pesos/liter (1.08 USD). The second, larger regime pools 57 observations, when the fuel price is more than 661.9 Chilean pesos/liter (1.08 USD).

Finally, for onion, the maximum $LM_{\gamma}$ statistics corresponds to a threshold value of 17.361. The first regime contains 76 observations and is active when the price of fuel is less than 571.1 Chilean pesos/liter (0.93 USD). The second, larger regime pools 98 observations, when the fuel price is more than 571.1 Chilean pesos/liter (0.93 USD).

Moreover, our results indicate that price transmission elasticity varies between models and regimes. The coefficients corresponding to the speed of adjustment to the long run equilibrium relationship has the correct negative sign and is statistically significant for all cases. The TVEC models estimates are presented in Table 2.

Table 2. Estimates for the regime dependent VECM

<table>
<thead>
<tr>
<th></th>
<th>Maize (MAI)</th>
<th>Tomato (TOM)</th>
<th>Onion (ONI)</th>
<th>Carrot (CAR)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Regime 1</td>
<td>Regime 2</td>
<td>Regime 1</td>
<td>Regime 2</td>
<td>Regime 1</td>
</tr>
<tr>
<td>Number of observations</td>
<td>79</td>
<td>98</td>
<td>101</td>
<td>73</td>
</tr>
<tr>
<td>Maximum LM statistic</td>
<td>17.834</td>
<td>17.877</td>
<td>17.834</td>
<td>17.877</td>
</tr>
<tr>
<td>Threshold limits</td>
<td>561</td>
<td>625</td>
<td>661.9</td>
<td>661.9</td>
</tr>
<tr>
<td>Price Transmission Elasticity</td>
<td>0.473</td>
<td>0.150</td>
<td>0.218</td>
<td>-0.089</td>
</tr>
<tr>
<td>Speed of Adjustment</td>
<td>-0.091</td>
<td>-0.737</td>
<td>-0.200</td>
<td>-0.377</td>
</tr>
<tr>
<td>ADF test (p-value)</td>
<td>(-0.05)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
</tr>
<tr>
<td>Mean of price difference (standard error)</td>
<td>0.046</td>
<td>0.631</td>
<td>0.012</td>
<td>0.005</td>
</tr>
<tr>
<td>(0.132)</td>
<td>(0.225)</td>
<td>(0.065)</td>
<td>(0.189)</td>
<td>(0.170)</td>
</tr>
<tr>
<td>Likelihood ratio test (p-Value)</td>
<td>43.665</td>
<td>55.332</td>
<td>38.768</td>
<td>39.556</td>
</tr>
<tr>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
<td>(&lt;0.01)</td>
</tr>
</tbody>
</table>
6. Discussion

Our initial hypothesis of a regime-dependent model in which the long-run relationship and the adjustment process displays threshold cointegration behaviour is supported by the results of the likelihood-ratio test (see Table 1). The speed of adjustment coefficients are significant for all variables, thus reinforcing the hypothesis of a causality pattern from Santiago market to regional market. These findings support the idea that independently of the quantity produced in Talca, the major effect of price adjustment is a result of the high demand and market concentration of the Central wholesale market than the favourable productive conditions in Talca. In this sense, results obtained for the full dataset suggest that for most of the products the price relationship is unidirectional and that the regional price only error corrects whereas the Santiago market dominates the price.

With respect to the Gonzalo and Pitarakis (2006) exogenous threshold driven cointegration confirming non-linear results. As previously identified it was detected two trade-share dependent long-run price elasticity relationships, we assume that the differences between each regime mainly depends of the product traded and the fuel prices.

At individual level, for maize, the larger regime with 98 observations, price transmission elasticity of 0.150 and speed of adjustment of -0.737 could be characterized the harvest regime when on average 55% of the maize offered on the Santiago correspond of product originated from Talca (March/April). The smaller regime probably considers the origin from another producer regions, such as, RM, VIII and IX regions. When compared with rest of products, the effect of the threshold variable is the highest for all variables, that is, lower price of fuel prices mostly affects the margin in which the trade between markets is profitable. It confirms our assumption that the reaction on arbitrage for annual crops, with a relative low producer prices, is much more elastic than cash crops that are characterized by a higher price segmentation.

Probably, tomato is one of the horticultural products with the more stable offer in the Chilean market. It is a vegetable of warm season and despite it high sensibility to the winter frost, the fresh tomato are present in the Chilean market during all seasons of the year (Ruiz-Machuca, 2015). The wide range of growing conditions offered by the country allows its cultivation from Arica and Parinacota Regions (North of Chile) to the Los Lagos Region (South of Chile). This coupled with the extent of crop varieties with indeterminate habit, enable market supply for much of the year on the central and regional vegetable wholesale markets. Tomato represents 70% of
all greenhouse production in Chile and it complements the outdoor production as soon as it is unable to supply the market (ODEPA, 2014). For tomato, the larger regime with 101 observations shows a price transmission elasticity of 0.218 and a speed of adjustment of -0.200. On the contrary, the smaller regime shows a price transmission elasticity of 0.089 and a speed of adjustment of -0.377. With this in mind, since the price transmission elasticities are trade dependent (Bakucs et al. 2014) and the seasonality effect is smoothed by the longer supply period. The differentes on the adjustment parameters appers to be a consequence of infraestructure or logistic than market supply/demand factors. Actually, during period april-september, most of the tomato offered in Santiago comes from regional markets but our results suggest that trader's profitability mainly depends of the speed in which the product are supplied to market and not of the fuel o price differences between regional and central markets.

The onion is the second vegetable in acreage and value of production in Chile. During last decade, the cultivated area has been around 10,000 hectares per year, with about 60% dedicated to "storage" onions and 40% for the early and mid-season onions. The main producing areas are located in the Central regions (Metropolitan, Fifth and Sixth regions). Onion presents the largest differences in the number of observations between the larger regime (117 observations) and smaller regime (57 observations). In the first, it shows a price transmission elasticity of 0.010 and a speed of adjustment of -0.065, while in the second, the price transmission elasticity is 0.122 and the speed of adjustment is -0.029. It presents the higher value of threshold exogenous variable among all varibles with a fuel prices of 661.9 Chilean pesos (1.08 USD). The characterization of the larger regime corresponds to the market conditions in January/March when the onion for storage are supplied to the wholesale markets. This is the regime in which the price relationship is unidirectional with the Santiago price not error correcting and thus dominating the Talca price. In contrast, the Talca price error corrects at relatively high speed. This may be attributed to the harvest season starting earlier in the central regions of Chile, implying that new onions are first sold on the Santiago market. Consequently, the initial price level for the new harvest is set on the Santiago market and is transmitted to the regional markets allowing relatively high price differences that give margin to profitable arbitrage opportunities.

Different to the previous variables, 70% of the carrot are cultivated on the Chilean central region, it has two planting season (March-May and October-November) and the post-harvest life is very short (3 weeks in average) (McKee et al. 2011). This situation generates a relatively low level of profitability with relation to another species, since the arbitrage advantages from seasonality
effect are more difficult to obtain. The larger regime has 98 observations and shows a price transmission elasticity of 0.149 with a speed of adjustment of -0.376. The smaller regime shows a price transmission elasticity of 0.473 and a speed of adjustment of -0.016. In this sense, it is not surprising the similar number of observations between regimes and the sensitivity of fuel prices as exogenous threshold. Another interesting statistic differentiating the two regimes is the mean of price difference and standard deviation. This statistic equals 0.046 in the first regime, but 0.631 in the larger regime suggesting higher volatility of traded quantities during harvesting season. Following Rodriguez and Hernandez (2005), this effect affects directly the gross margin obtained by the wholesale markets since the irregular supply behavior and low post-harvest period of carrot. From our results, we assume a transition effect to carrot producers which also affect the planting decision and thus, modified their productive structure at farm level. However, this conclusion is beyond the scope of our article.

7. Conclusions

In this paper, we have analyzed the dynamics of spatial horticultural price transmission between a central wholesale market (Santiago) and a regional wholesale market (Talca), both located within a distance of approximately 280 km. The objective is to identify the pathways of seasonality effects, asymmetries on the price transmission and the role played by the Chilean fuel prices in horticultural price market integration by the inclusion of an exogenous threshold variable. We propose a novel procedure to estimate a regime-dependent VEC model accounting for threshold effects not only in the short-run adjustment towards the long-run equilibrium but in the long-run equilibrium relationship as well. This method seems to be particularly suitable in settings of irregular seasonal price transmission, typical for horticultural products.

We found clear evidence of threshold cointegration with Chilean retail fuel prices serving as the threshold variable. For all variables (maize, tomato, onion and carrot) we identify two price transmission regimes which are characterized by different equilibrium relationships as well as short-run adjustment processes towards this equilibrium. Our econometric results fit well with the actions on the central and regional horticultural markets. This research could be extended to drilling down the effect of specific public regulations on the degree of market integration among vegetable markets. For example, since fuel price appears to be an important variable to take account for the farmer's planting decision, the implementation of
policies oriented to increase the efficiency of the price transmission between markets could affect positively the gross margin from producers and wholesale market on the commercialization chain.

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8. Bibliographic references


