

Spatial and cross-product price linkages in the Brazilian pine timber markets

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ARTICLE INFO

Keywords:

Timber markets
Price transmission
Timberland investments
STAR
Market linkage

ABSTRACT

The South of Brazil is one of the most attractive regions for timberland investments in the world. High productivity and relatively attractive timber prices have gained attention from timberland investors. However, as in most emerging countries, it is not very clear how prices are transmitted across products and markets. Having this information is essential to strategic planning as well as understand the market structure. We investigate market linkages of the stumpage price of five products (fuelwood, pulpwood, sawtimber, veneer, and special veneer) in the three main pine producing states in Brazil (Paraná, Santa Catarina and Rio Grande do Sul). We use linear and regime shifting models and check the effect of external shocks on price transmission. The nonlinear process is observed main on high-grade timber (veneer and special veneer), possibly driven by their price recovery after the recession. Our results show that the spatial and between product price ratio converges back the market equilibrium within 10 months in the pine stumpage market in Brazil. This outcome indicates this market is efficient with small opportunities for arbitrage profits.

1. Introduction

Institutional investors have increased their interest in timber investments because of the relatively good return, low financial risk, and low correlation with other financial sectors (Redmond and Cubbage, 1988; Sun and Zhang, 2001; Wan et al., 2015; Zhang and Mei, 2019). Although the economic and political stability has often attracted timberland investors to the United States and Europe, the relative low returns in recent years (Cubbage et al., 2020) and the desire for further diversification of timberland portfolio (Busby et al., 2020) have driven investment to other regions, especially in Latin America. In 2017, the investable timberland assets in Latin America totaled \$55 billion, 63% in Brazil and the remaining 37% in other countries like Chile, Costa Rica and Uruguay (Newforests, 2017).

A large share of the timberland financial returns are driven by biological growth and timber prices, often termed the “biological beta” (Caulfield, 1998; Conroy and Miles, 1989; Mei et al., 2015). Large investments in silviculture and genetics have substantially increased the biological growth of many tree species and provided a good understanding of silvicultural treatments and their impact on forest structure and financial returns (D’Amato et al., 2017). Timber prices, on the other hand, are more volatile and harder to predict because they are usually determined by a complex web of multiple interactions among industries, landowners, and government agencies. In the face of

uncertainty, price transmission studies are key in the understanding of forest products markets and price projections, especially those located in developing countries. In this article, we aim to investigate how prices of pine timber products are interrelated across the Brazil’s southern region, where timber prices have been shown to substantially impact the return-risk relationship in timberland investments.

In timberland investment portfolios, for example, pulpwood price fluctuations of Eucalyptus plantations in Brazil can change the internal rate of return from 2.35% to 12.76%, while chip-n-saw prices of Pine plantations can alter from 13.41% to 15.05% (Chudy et al., 2020). By using time-varying smooth transition autoregressive models, we evaluate the spatial and cross-product price linkages of pine timber products in the South of Brazil. An understanding of price movements and their spatial interconnections can help timberland investors in making investment decisions in this important timber-producing region of the world.

Brazil is particularly important in timber production and forest products processing. Its share of the world’s total industrial wood production is estimated at 7%; representing 23% of the global export of pulp and paper and 2% of the export of solid wood products (Food and Agriculture Organization of the United Nations, 2018). In addition, the fast growth rates of Brazil’s planted forests provide financial advantages based on both efficiency and shorter rotations. They also can reduce physical risks, such as wildfire and pests, as well as political and

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<https://doi.org/10.1016/j.forpol.2020.102186>

Received 30 April 2019; Received in revised form 19 April 2020; Accepted 22 April 2020

Available online 16 May 2020

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macroeconomic risks that are characteristic of developing markets. These characteristics have profound impacts on the allocation of timberland portfolio; according to Busby et al. (2020), timberland investors could see an 8.5% of financial return by allocating 40% of their portfolio in southern Brazilian plantation, the highest among 14 possible combinations of species and regions to invest.

In 2017, Timberland Investment Management Organizations (TIMOs) owned and/or managed 1.09 million hectares in Brazil (Kanieski da Silva et al., 2017). However, the lack of publicly available information on Brazilian timber markets may discourage risk-averse investors. Since timber prices convey information about timber scarcity, understanding stumpage prices dynamics is essential to set investor expectations and inform private and decision makers.

1.1. Price transmission and market dynamics

Price transmission studies can reveal market conditions, like market competition (Sun and Ning, 2014) and arbitrage behavior (Liefert and Persaud, 2009), that might distort producer and consumer decisions, with consequences to economic welfare (Cudjoe et al., 2010). One essential feature discussed for decades by economists about spatial market equilibrium is the concept of the Law of One Price (LOP). The LOP postulates that, in an efficient market, the price difference of identical goods produced in different locations should be equal to their transaction costs (Fackler and Goodwin, 2001). Violating the LOP would lead to arbitrage profits, where profit-seekers maximize their returns by moving goods from low-priced to high-priced regions. The ultimate consequence is suboptimal total welfare since this movement might indicate arbitrage profits, non-competitive practices, and thus the presence of deadweight losses (O'Connell and Wei, 2002).

A large body of literature has investigated LOP and price dynamics in the forest sector. Nanang, 2000, for instance, rejected LOP in the lumber markets among five Canadian provinces and suggested these markets could be modeled separately. Størdal and Nyrud (2003) examined market efficiency by testing market cointegration between domestic, imported and exported sawlogs in Norway. Among other policy implications, this study revealed that the Norwegian sawmilling industry was not importing sawlogs to curb domestic demand, and highlighted the weak impact of international prices in the local market. Investigating market integration between timber products at different stages of manufacturing, Zhou and Buongiorno (2005) showed that, even though standard statistical tests did not show any cointegration, some markets such as sawlog and lumber could affect each other asymmetrically. According to these authors, policies (such as planting incentives) focused on sawlog have little impact on lumber prices, whereas lumber market policies (such as household subsidies) have a large impact on sawlog prices.

Another significant number of studies have focused on LOP and price cointegration of stumpage prices; their results vary according to location and period (Bingham et al., 2003; Mei et al., 2010; Nagubadi et al., 2001; Prestemon, 2003; Prestemon and Holmes, 2000; Yin et al., 2002). Collectively, these studies show a complex relationship between several inter-regional markets driven by local market forces, such as timber inventory, industry demand, and harvest volumes.

Most recently, testing vertical and spatial price transmission in various forest product markets advanced from a linear cointegration structure to a non-linear approach. Goodwin et al. (2011) and Hood and Dorfman (2015) developed multiple specifications of smooth transition autoregressive (STAR) models to examine market linkages in oriented strand board (OSB) and pine stumpage markets, respectively. Their analyses revealed the spatial market linkage in both products predicted on spatial economic theory. Similarly, Parajuli and Chang (2015) and Parajuli and Zhang (2016) applied the Johansen et al. (2000) modified cointegration tests in the presence of structural breaks to estimate vector error correction models of softwood stumpage and lumber markets, respectively. Lastly, Chudy et al. (2020) found very little

evidences of cointegration between roundwood prices of different countries.

Our study complements and expands earlier investigation in several ways. Investigation about timber price dynamics focused mostly in developing countries, where transportation systems, access to information, and regulatory institutions are more advanced than in developing countries. Spatial price transmission, for instance, strongly dependent the distance between markets, and therefore on transportation costs (Bingham et al., 2003). Brazil is ranked 56th among 160 countries (behind Mexico, Chile, and South Africa) in logistics infrastructure, according to the Logistic Performance Index (LPI) (World Bank, 2018). This implies a great restriction on regional and international trade. Also, cumbersome tax requirements might lead to poor economic outcomes. For instance, Brazilian companies have to pay extra taxes (locally known as ICMS or *Imposto sobre Circulação de Mercadorias e Serviços*) if their products cross over different state borders. The ICMS discourages inter-regional trade, hinders transportation goods to ports, and impacts price linkage across states.

This article specifically evaluates the cross-product and spatial price transmissions between various pine products in the Brazilian markets. Clarifying the long-term and time-varying relationship between stumpage prices across different regions and products provide valuable information to both private investors and public institutions. We evaluate how timber markets (specifically pine stumpage prices) have been interconnected over the years across three states in the South of Brazil – Paraná, Santa Catarina, and Rio Grande do Sul – and five timber products – (1) fuelwood, (2) processing, (3) sawtimber, (4) veneer and (5) special veneer.

The structure of this paper is the following. First, we present a brief description of the Brazilian stumpage markets, followed by the concept of market linkages and the econometric models. Then, we present the results and discussion, and conclude with possible future studies.

2. Brazilian timber markets

The total forest plantation area in Brazil is 7.8 million hectares (IBA – Indústria Brasileira de Florestas, 2019), of which Eucalyptus plantation occupies 5.7 million hectares and Pine plantation covers 1.6 million hectares.¹ Eucalyptus plantations are mostly planted in the Central West and Southeast regions (responsible for 57% of total Eucalyptus plantations in the country), while Pine plantations are predominantly located in the South region (88% of the total area). The Brazilian planted forest markets are dominated by vertically-integrated and capital-intensive companies (IBA – Indústria Brasileira de Florestas, 2019). The concentration of land ownership in these vertically-integrated companies makes investing in Eucalyptus less attractive to timberland investors.²

On the other hand, the Pine timber markets present a more competitive structure with a greater number of market players both in demand and supply. These markets are mostly located in the South of Brazil, the states of Santa Catarina (SC), Paraná (PR) and Rio Grande do Sul (RS) (Figs. 1 and 2). Despite evidence of some market concentration associated with high transportation costs and spatial economics (Rogers and Sexton, 1994), the South of Brazil has a relatively competitive timber market similar to the one in the U.S. South and Scandinavia.

According to the Brazilian Institute of Geography and Statistics (IBGE, 2018), the South of Brazil produced altogether 94% of the country's total pine timber output in 2017; PR produced 46% of the total output, whereas SC and RS produced 38% and 10% respectively (Fig. 1A). The main timber buyers are located in clusters in the center of

¹ The remaining 0.5 million hectares are composed by other species, such as teak.

² Similar cases have been observed in the US South regarding market power of the mills (e.g., Mei and Sun 2008, Silva et al. 2019).

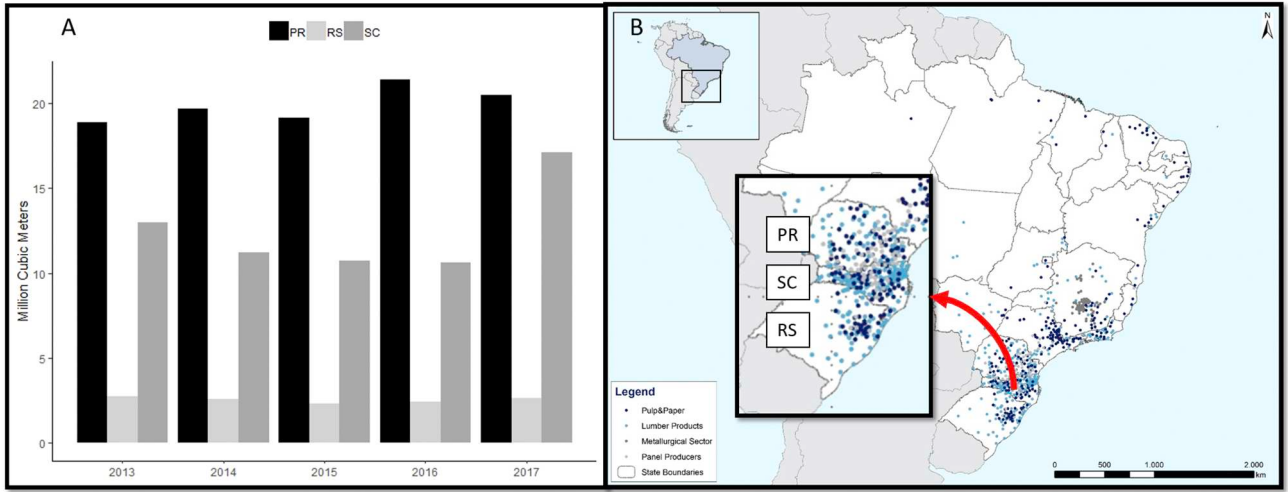


Fig. 1. (A) Pine timber Production in the South of Brazil per State. (B) Timber consumers in Brazil by industry type. PR: Paraná, SC: Santa Catarina and RS: Rio Grande do Sul. Sources: (A) Brazilian Institute of Geography and Statistics (IBGE, 2018). (B) Pöyry Global 2018.

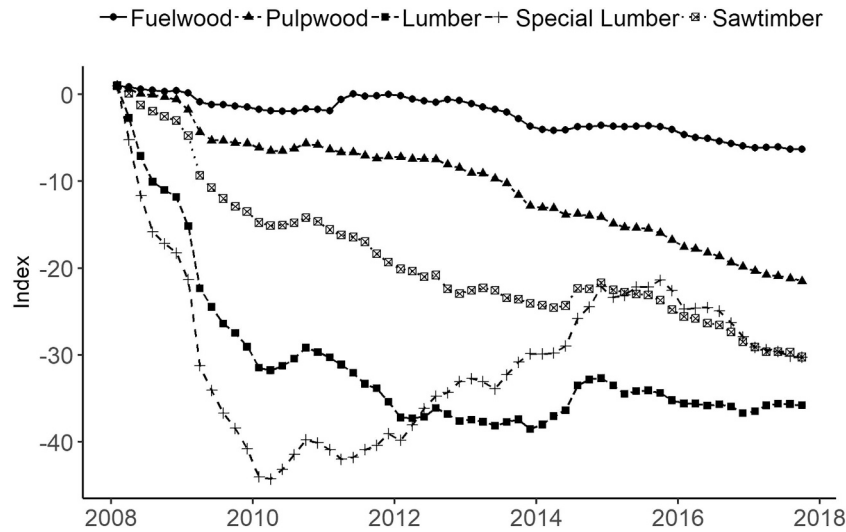


Fig. 2. Average deflated timber prices by product indexed. Sources: (A) Brazilian Institute of Geography and Statistics (IBGE, 2018). (B) STCP Engenharia - 2018.

PR, western and eastern SC, and northeastern RS (Fig. 1B). These mills contributed to \$939.6 million to the local economy only in wood transactions (30% of the total in Brazil) and are an important source of jobs and taxes in the southern Brazilian states.

3. The economic model

We develop an economic model to examine linkages among pine timber markets in southern Brazil for five products pine stumpage markets. The models relied on Smooth Transition Autoregressive (STAR) estimation and analyzed both spatial stumpage price markets and cross-products markets.

3.1. Spatial price linkage

We define price linkage as a measure of the degree to which demand and supply shocks in one region are transmitted to another (Fackler and Goodwin, 2001). Mathematically this spatial relationship can be described as:

$$R_{AB} = \frac{dp_B/ds_A}{dp_A/ds_A} \quad (1)$$

where $p_{A(B)}$ is the price of a good in market A (B), and R_{AB} it the price transmission ratio between regions A and B, which equals one if markets A and B are perfectly integrated and zero if there is no integration at all. Notice that external shocks do not depend on trade between regions A and B. In fact, markets can be cointegrated because of other economic factors such as supply, demand and industry concentration (Bingham et al., 2003; Prestemon and Holmes, 2000). In the case of stumpage prices, spatial or cross-product cointegration might also exist because of changes in the supply and demand of their derived products, such as lumber (Ning and Sun, 2014).

3.2. Cross-product price linkage

Cross-product price linkage can also be defined according to Eq. 1, but instead of multiple regions, there are multiple products under different stages within the supply chain. Timber is a special case because the biological growth makes trees larger and suitable for different uses, from lower to higher grades, e.g. from pulpwood to veneer. Landowners will, therefore, postpone harvest if they expect a change in prices greater than the discounted storage and capital costs. Washburn and Binkley (1990) expressed this intertemporal arbitrage condition for the stumpage prices as:

$$E_t[P_{t+1}] - P_t = r_t - (g_t - c_t) \quad (2)$$

where P_t is the stumpage price during period t , g_t is the biological growth rate, c_t is the cost of storing the marginal unit and r_t is the capital cost of holding stumpage during period t or the rate of return required by investors to stock timber.

Assuming that the difference between the cost of storing the marginal unit (c_t) offsets the growth rate (g_t), if the expected increase in stumpage prices ($E_t[P_{t+1}] - P_t$) is greater than the marginal capital cost (r_t), then landowners will withhold stumpage from the market during the current period and vice-versa. This dynamic is also observed cross-product like in $E_t[P_{t+1}^{k_1}] - P_t^{k_2} \geq r_t$, where k_2 and k_1 are the higher and lower timber grade respectively. Landowners might opt to thin their forest instead of clearcutting to enjoy more attractive prices of sawtimber and/or veneer. Moreover, at the end of the timber production cycle landowner use similar logic to decide whether to replant or change the land use.

This timber supply dynamic links different timber products, i.e. pulpwood and sawtimber. As the expected sawtimber stumpage price increases, more pulpwood is withheld, driving current pulpwood stumpage prices up, and vice-versa. The equilibrium between future and present stumpage prices will hold as long as the forest inventory is not exhausted and changes in supply and demand occur smoothly (Washburn and Binkley, 1990).

4. Data

We analyzed spatial and cross-products linkages of bimonthly stumpage prices series, 2008 to 2017, of five product classes in three states in the South of Brazil (Paraná, Santa Catarina, and the Rio Grande do Sul). The timber grade of each product based on its top-end log diameter in centimeters is (1) Fuelwood (< 8 cm), (2) Pulpwood (8–15 cm), (3) Sawtimber (15–25 cm), (4) Veneer (25–35 cm) and, (5) Special Veneer (> 35 cm).

In every product described, there was a substantial drop in prices since the economic recession in 2008. The largest grades have suffered a higher fall; between 2008 and 2010, special veneer and veneer had a sudden drop of 44% and 38% respectively and after 2010 they seem to follow a positive trend. On the other hand, fuelwood, pulpwood, and sawtimber prices have been decreasing consistently since 2008. Fuelwood prices shrank 6% and sawtimber 32% between 2008 and 2017. Most of this change is linked to the market structure, while higher grades are mostly exported, the lower ones are consumed locally.

5. Econometrics

Cross products and spatial market linkages are based on possible random movements of two or more price series in the short run. However, in the long-run, economic forces prevent them from moving too far apart (Fackler and Goodwin, 2001). To investigate the price linkage across different markets and products, we used a Smooth Transition Autoregressive Model (STAR) (Terasvirta, 1994) similar to Goodwin et al. (2011), Hood and Dorfman (2015) and O'Connell and Wei (2002).

Consider two nonstationary log prices located in markets i and j defined as $\ln(p_i)$ and $\ln(p_j)$. Given the price ratio $y = \ln(p_i/p_j)$ and the transaction cost per unit l ($0 \leq l \leq 1$), here composed only by transportation costs for simplicity, take the Samuelson 'iceberg form', only the fraction $(1 - l)$ of a good arrives when shipped to other markets. This framework creates a band for the relative prices, $-\ln(1 - l) \geq y \geq \ln(1 - l)$ where arbitrage activities are not possible. The relative prices (y) could, therefore, move randomly, but when reaching an extreme value, transaction costs would prevent prices to move from outside the band.

A similar concept could be applied to the relative prices of different

products in the same market. Assume the prices of the two timber products are p_{k2} and p_{k1} , where $p_{k2} \geq p_{k1}$ - product k_2 has a higher grade than k_1 (e.g., sawtimber and pulpwood). The price band is then defined as $-\ln(1 - b) \geq \ln(p_{k2}/p_{k1}) \geq 0$, where b ranges from 0 to 1, and could be interpreted partially as the expected opportunity cost per unit. This cost includes silvicultural operations and the forgone opportunity of harvesting lower grades for harvesting future larger ones given current market information. The right-hand side equals zero because higher grades could substitute lower grade products (not the inverse because of technological limitations); in fact, timber buyers would be indifferent between lower and higher grades if their price ratio approximates zero.

We modeled the price ratio between $\ln(p_{i1})$ and $\ln(p_{j1})$ using different formulations. We start by testing if their linear relationship (e.g., $y_t = \ln(p_{i1}/p_{j1})$) is stationary.³ Thus, an Error Correction Model (ECM) in the same fashion as the Augmented Dickey-Fuller (ADF) test, can be used to model the long-term relationship between two price series as shown in Eq. 3:

$$\Delta y_t = \alpha_0 + \sum_{p=1}^P \delta_p \Delta y_{t-p} + \beta y_{t-1} + e_t \quad (3)$$

where $\Delta y_t = y_t - y_{t-1}$, δ_p and β are the parameters to be estimated, and p is the optimal lag calculated by the Akaike Information Criterion (AIC). If $\beta < 0$, y_t is stationary and $\ln(p_i)$ and $\ln(p_j)$ are cointegrated.

Eq. 3 assumes transaction costs are constant and proportional to commodity prices and are inconsistent with discontinuous trade (Baulch, 1997). Also, adjustment to the equilibrium of price parities is often non-linear because of the high transaction costs (Dumas, 1992).⁴ Relative prices might have periods in which there is a strong cointegration and others that there is no linkage between them at all. We evaluated these dynamics by expanding the specification of Eq. 3 to a STAR framework that considers a change between regimes from stationary and non-stationary as shown in Eq. 4:

$$\Delta y_t = \left(\alpha'_0 + \sum_{p=1}^P \delta'_p \Delta y_{t-p} + \beta' y_{t-1} \right) \underbrace{G(s_t; \gamma(\eta), c)}_{\text{Regime One}} + \left(\alpha_0 + \sum_{p=1}^P \delta_p \Delta y_{t-p} + \beta y_{t-1} \right) \underbrace{G(s_t; \gamma(\eta), c)}_{\text{Regime Two}} + e_t \quad (4)$$

where β (β') < 0 are tested, $G(s_t; \gamma(\eta), c)$ is the transition function and value ranges from zero to one. s_t is the transition variable that determines the threshold for regime shifting, $\gamma(\eta)$ is the speed in which the relative prices switch between regimes, the higher the value the faster is the change between regimes. Because γ must be positive, we opted to transform $\gamma(\eta) = -\exp(-\eta)$, thus ensuring positive values without imposing further constraints in the model (Goodwin et al., 2011). c is the threshold value and e_t is the error term. The representation of $G(s_t; \gamma(\eta), c)$ here is a logistic (Eq. 5) or exponential (Eq. 6) functions defined as:

$$G(s_t; \gamma, c) = \left[1 + \exp \left\{ \frac{-\gamma(\eta)(s_t - c)}{\sigma_{st}^2} \right\} \right]^{-1} \quad (5)$$

³ ADF tests cointegration of a log price ratio, $y_t = \ln(p_{i1}/p_{j1})$ assuming efficient markets request that the relation $\ln p_{it} = \beta_0 + \beta_1 \ln p_{jt}$ is restricted by $\beta_0 = 0$ and $\beta_1 = 1$. Similar method is employed in Goodwin et al. (2011) and Hood and Dorfman (2015). The individual ADF test for stationary fails to reject the non-stationarity for every product and state; these results are available under request to the authors.

⁴ Harvesting costs are embedded in the transaction costs when studying stumpage price parity.

Table 1
Selected pairs by the Logistic Smooth Transition Autoregressive Model (LSTAR).

	PR	RS	SC		$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{PR}/p_t^{SC})$	$\ln(p_t^{RS}/p_t^{SC})$
$\ln(p_t^E/p_t^{PW})$	No	No	No	Energy (< 8 cm)	Yes	No	Yes
$\ln(p_t^E/p_t^V)$	Yes	No	Yes	Pulpwood (8–15 cm)	No	No	Yes
$\ln(p_t^E/p_t^{SV})$	No	No	No	Sawtimber (15–25 cm)	Yes	Yes	No
$\ln(p_t^E/p_t^{ST})$	No	Yes	No	Veneer (25–35 cm)	No	No	Yes
$\ln(p_t^{PW}/p_t^V)$	No	No	No	Special Veneer (>35 cm)	No	No	Yes
$\ln(p_t^{PW}/p_t^{SV})$	No	No	No				
$\ln(p_t^{PW}/p_t^{ST})$	No	No	Yes				
$\ln(p_t^V/p_t^{SC})$	No	No	No				
$\ln(p_t^V/p_t^{ST})$	No	No	No				
$\ln(p_t^{SV}/p_t^{ST})$	No	No	Yes				

Note: PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. E = Energy (< 8 cm), PW = Pulpwood (8–15 cm), ST = Sawtimber (15–25 cm), V = Veneer (25–35 cm), SV = Special Veneer (> 35 cm).

$$G(s_t; \gamma, c) = 1 - \exp\left(-\frac{\gamma(\eta)(s_t - c)^2}{\sigma_{st}^2}\right) \gamma > 0 \quad (6)$$

The functions in Eq. (5) and (6) embedded into Eq. (4) are called the LSTAR and ESTAR models respectively (Chan and Tong, 1986). We selected between LSTAR and ESTAR by carefully analyzing their coefficients and model performance (AIC, BIC, etc...).⁵ Once $G(s_t; \gamma(\eta), c)$ is defined, we specify the transition variable s_t as the six periods (12 months since the data is bi-monthly) moving average of the log of prices (y_t) (Eq. 7).

The economic rationality for six periods is that the movement between regimes should be affected by a shock large enough to change the relative price patterns within a year. In addition, timber supply, demand, and inventory are commonly estimated at an annual basis, thus, it is a reasonable assumption that it takes at least a year for every economic agent to have the information about market changes.

$$s_t = 1/6 \sum_{n=1}^{N=6} y_{t-n} \quad (7)$$

5.1. Model specification

A key component of studying price transmission across different regions and products is to test unit root and linearity. We follow five steps to define the final models for each pair of prices: (i) we tested unit root against stationarity and nonlinearity in the same fashion as in Goodwin et al. (2011), (ii) If the null hypothesis in the first step is rejected, we tested a Logistic STAR against an Exponential STAR model, (iii) alternatively, if the test failed to reject the null hypothesis in step (i) we tested for unit root against linear stationarity using the ADF test, (iv) if the test fails to reject unit root in the step (iii) we tested linearity on the AR model and, lastly (v) if linearity is rejected on step (iv), we tested a Logistic against Exponential AR - STAR model.

5.1.1. Testing unit root and nonlinearity framework

The nonlinear dynamics of economic variables (e.g., price ratio, purchasing power parity (PPP), etc.) have increased the concern about the persistent failure of the standard ADF test to reject the null of a unit root. Over the last decades, economists have investigated alternative unit root tests that account for the nonlinear characteristics. The joint analysis of nonstationary and nonlinearity in the context of threshold cointegration was first popularized by Balke and Fomby (1997). Kapetanios et al. (2003) extended Balke and Fomby works by developing a unit root test against the ESTAR model. Goodwin et al. (2011) used a similar approach proposed by Eklund (2003) to test the linear

unit root model against the Generalized ESTAR model. The procedures proposed by the literature test the AR(p) model

$$\Delta y_t = \alpha_0 + \sum_{p=1}^P \delta_p \Delta y_{t-p} + e_t \quad (8)$$

against an ESTAR or LSTAR model as in the Eq. (4).

Intuitively, the linearity hypothesis is rejected when $H_0 : \gamma = 0$. Luukkonen et al. (1988), suggested that the $G(s_t; \gamma, c)$ could be extended to a suitable Taylor series approximation. For the LSTAR and ESTAR model proposed in Eq. (5) and (6), a third-order Taylor approximation would be appropriated (Terasvirta, 1994):

$$\Delta y_t = \alpha'_0 + \sum_{p=1}^P \delta'_p \Delta y_{t-p} + \beta_0 y_{t-1} + \sum_{i=1}^3 \omega_i s_t^i + \sum_{i=1}^3 \sum_{p=1}^P \vartheta_i (\Delta y_{t-p} s_t^i) + \sum_{i=1}^3 \beta_i (y_{t-1} s_t^i) + e_t \quad (9)$$

The test imposes, therefore, the restriction $\beta_0 = \omega_1 = \dots = \vartheta_1 = \dots = \beta_1 = \dots = 0$ on Eq. (9). We refer the null hypothesis and F-statistic as H_{0lur} and F_{lur} , which in turn has $(7 + 2p)$ and $T - (8 + 4p)$ degrees of freedom. The peculiarity in this approach is that the F-Statistic is no longer associated with the usual limiting distribution under the null of linearity and a unit root. After estimating the F_{lur} , we then use a dynamic bootstrap of the null model's estimated residuals to construct an empirical distribution of the F_{lur} and extract the p -value (Balagtas and Holt, 2009).

5.1.2. Testing logistic STAR and exponential STAR

If the null hypothesis is rejected against a nonlinear stationary series on step one, we then use a similar approach to select between LSTAR and ESTAR. According to Terasvirta (1994), the Taylor series approximation of an LSTAR and ESTAR models are represented in Eq. (9) and (10) respectively.

$$\Delta y_t = \alpha'_0 + \sum_{p=1}^P \delta'_p \Delta y_{t-p} + \beta_0 y_{t-1} + \sum_{i=1}^2 \omega_i s_t^i + \sum_{i=1}^2 \sum_{p=1}^P \vartheta_i (\Delta y_{t-p} s_t^i) + \sum_{i=1}^2 \beta_i (y_{t-1} s_t^i) + e_t \quad (10)$$

The author suggested a series of F-tests to specify choose between the logistic and exponential transition functions. First, they suggest testing between Eq. (8) and (9), using a simple F-test, $H_{01} : \omega_3 = \dots = \vartheta_3 = \dots = \beta_3 = \dots = 0$. However, only rejecting H_{01} would not necessarily reject the ESTAR model in special cases.⁶ Therefore two other F-test are suggested like $H_{02} : \omega_2 = \dots = \vartheta_2 = \dots = \beta_2 = \dots = 0 \mid \omega_3 = \dots = \vartheta_3 = \dots = \beta_3 = \dots = 0$ and $H_{03} : \omega_1 = \dots = \vartheta_1 = \dots = \beta_1 = \dots = 0 \mid \omega_2 = \dots = \vartheta_2 = \dots = \beta_2 = \dots = 0 \mid \omega_3 = \dots = \vartheta_3 = \dots = \beta_3 = \dots = 0$.

⁵ We also tried other transition functions like Quadratic STAR - $G(s_t; \gamma, c) = \left[1 + \exp\left\{\frac{-\gamma(\eta)(s_t - c)^2}{\sigma_{st}^2}\right\}\right]^{-1}$ - but the QSTAR specification had a poorer performance than LSTAR and ESTAR.

⁶ Terasvirta (1994) presented an extensive explanation about these steps. We omitted them because of the different scope of this paper.

Table 2
Augmented Dickey - Fuller (ADF) test for the Log-Ratio Stumpage Prices Series by Products.

PR				SC			RS		
	Statistic	Lags	Cointegrated?	Statistic	Lags	Cointegrated?	Statistic	Lags	Cointegrated?
$\ln(p_t^E/p_t^{PW})$	-0.84	1	No	-1.13	1	No	-0.61	1	No
$\ln(p_t^E/p_t^V)$	0.87	1	No	0.46	1	No	0.19	1	No
$\ln(p_t^E/p_t^{SV})$	0.93	1	No	0.87	1	No	1.25	1	No
$\ln(p_t^E/p_t^{ST})$	0.75	1	No	-0.38	3	No	-0.25	1	No
$\ln(p_t^{PW}/p_t^V)$	1.97	1	No	1.86	1	No	0.83	1	No
$\ln(p_t^{PW}/p_t^{SV})$	2.20	1	No	2.73	1	No	2.50	1	No
$\ln(p_t^{PW}/p_t^{ST})$	2.28	1	No	0.07	3	No	0.50	1	No
$\ln(p_t^V/p_t^{SV})$	0.44	1	No	1.73	1	No	0.98	1	No
$\ln(p_t^V/p_t^{ST})$	0.68	1	No	1.75	2	No	0.51	1	No
$\ln(p_t^{SV}/p_t^{ST})$	1.37	2	No	2.63	2	No	2.67	1	No

Note: PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. E = Energy (< 8 cm), PW = Pulpwood (8–15 cm), V = Veneer (15–25 cm), SV = Special Veneer (25–35 cm), ST = Sawtimber (> 35 cm). The optimal lag was determined by minimizing Akaike Information Criterion(AIC). * $p < .1$, ** $p < .05$, *** $p < .01$.

Table 3
Augmented Dickey - Fuller (ADF) test for the Log-Ratio Stumpage Prices Series by State.

$\ln(p_t^{PR}/p_t^{RS})$			$\ln(p_t^{PR}/p_t^{SC})$			$\ln(p_t^{RS}/p_t^{SC})$			
Statistic	Lags	Cointegrated?	Statistic	Lags	Cointegrated?	Statistic	Lags	Cointegrated?	
Energy (< 8 cm)	-0.83	1	No	-0.71	1	No	-1.90*	3	Yes
Pulpwood (8-15 cm)	-1.16	1	No	0.52	1	No	-1.68*	3	Yes
Sawtimber (15-25 cm)	0.47	2	No	-0.47	1	No	-1.50	2	No
Veneer (25-35 cm)	-1.40	1	No	-2.66***	1	Yes	-0.94	1	No
Special Veneer (> 35 cm)	0.55	1	No	-0.16	1	No	-1.29	1	No

Note: PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. The optimal lag was determined by minimizing Akaike Information Criterion(AIC). * $p < .1$, ** $p < .05$, *** $p < .01$.

Teräsvirta proposed the following decision system to select between ESTAR and STAR based on the hypothesis described above: after rejecting the general null hypothesis (item 5.1.1), carry out the three F tests (H_{01} , H_{02} and H_{03}). If the p-value of the test H_{02} is the smallest of the three, select the ESTAR model; if not, choose an LSTAR model. We also add one more restriction by requiring a smaller standard error of the error term in the selected model. Our initial analysis showed that in some cases the specification selected by the F-tests sequence presented poor fit.

5.1.3. AR(p) linear and non-linear (LSTAR and ESTAR)

If the data generating process of a price ratio has a unit root, we work with its first difference in an AR(p) model. Also, we similarly tested linearity as the unit root vs nonlinear stationary test proposed in section 5.1.1. Using Eq. (8), the linear hypothesis was defined as $H_0: \beta_0 = \omega_1 = \dots = \vartheta_1 = \dots = 0 \mid \beta_0 = \beta_1 = \dots = \beta_4 = 0$. The F-test distribution here is also derived from a bootstrap simulation; we named the F statistic as F_l for this test.

In case the linear hypothesis is rejected, we employed the similar steps of section 5.1.2 to choose between LSTAR and ESTAR but we conditioned the hypothesis to $\beta_0 = \beta_1 = \dots = \beta_4 = 0$.

5.2. Model estimation

The linear model is estimated by standard OLS methods, while the STAR specification used a conditional Non-Linear Least Square (NLS) (Franses and van Dijk, 2003). This method minimizes the sum of the squared error (SSR) by specifying a priori value of η , and estimated the remainder coefficients. We repeated these steps over a range of 120 a priori values of η (from -6 to 6, increasing by 0.1) and selected the model with the lowest AIC.

5.3. Generalized impulse response function

We employed a Generalized Impulse Response Function (GIRF) to study how external shocks on stumpage prices can affect their relationship through time (Koop et al., 1996). The GIRF can review permanent or transitory price adjustments trends depending on the nature of the data. Mathematically, the response for a specific shock δ and history Ω_{t-1} can be expressed as:

$$GIRF(h, \delta, \Omega_{t-1})_{\Delta y} = E[\Delta y_{t+h} | (e_t = \delta, \Omega_{t-1})] - E[\Delta y_{t+h} | (e_t = 0, \Omega_{t-1})] \quad (11)$$

where h is the forecasted time horizon, δ is the external shock at period $T + 1$, and Ω_{t-1} is the historical variable ($\Delta y_{t+1}, \dots, \Delta y_{t+p}$). In this study, the response of the price ratio is evaluated for a horizon of 24 months. We included positive and negative shocks at period $h = 0$, with a magnitude of one and two times the standard error of the residuals.

The estimation steps are as follows. We selected the histories available for each pair, and for each combination of history and shock, we calculated the $GIRF(h, \delta, \Omega_{t-1})_{\Delta y}$ for $h = 0, 2, \dots, 24$. Also, we obtained the expected values in Eq. 11 by randomly sample the model's residuals 999 times and recursively calculating the realization of the price ratio. The final impulse response functions are computed by summing those estimated by

$$GIRF(h, \delta, \Omega_{t-1}) = \sum_{i=0}^h GIRF(h, \delta, \Omega_{t-1})_{\Delta y} \quad (12)$$

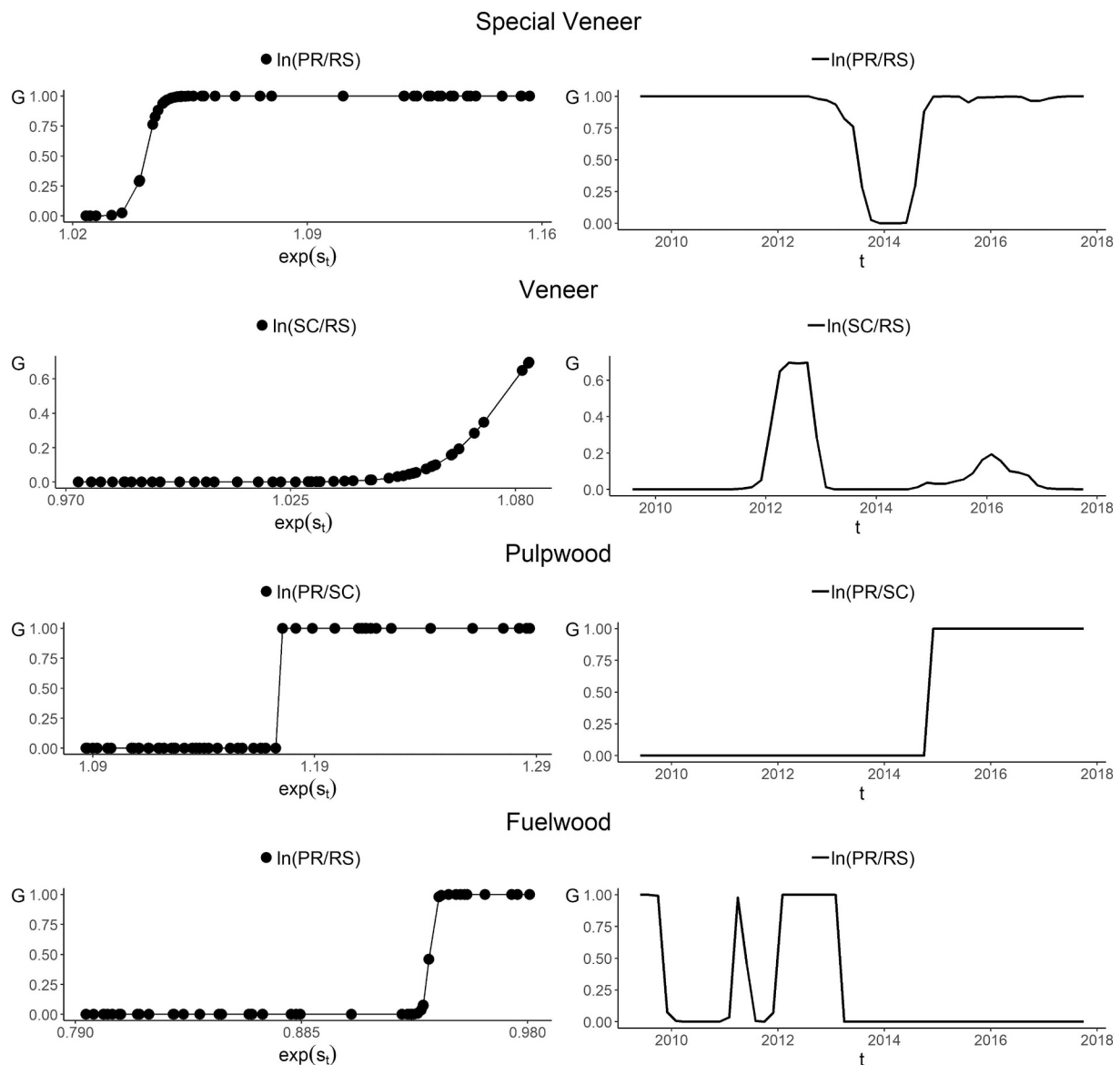


Fig. 3. Parana, SC and Rio Grande do Sul – Cross Products. Odd columns have the Transition Function versus the respective transition variable, while the even columns have the Transition Function over time (Feb/09 to Jun/17). p_t = stumpage price, E = Energy, PW = Pulpwood, V = Veneer, SV = Special Veneer, ST = Sawtimber.

6. Empirical results

6.1. Model specification

Tables 1 and 2 show the three tests for unit root and linearity. The unit root was rejected against a stationary nonlinear model in a few pairs. Stationary and nonlinear specifications are selected for the spatial price ratio of special veneer and fuelwood between PR and RS, veneer between SC and RS, and pulpwood between PR and SC. For cross-product, the price ratio of veneer and fuelwood in PR, special veneer and veneer in RS and special veneer and sawtimber in SC.

The ADF test indicated that only one price pair (sawtimber/fuelwood in RS) rejected the unit root against linear stationary. Therefore, the adequate specification of most of the pairs was autoregressive, with exception of veneer/sawtimber in RS which in turn presented

autoregressive and nonlinear specification. Among the nonlinear models, their best specification was an LSTAR model (Table 3) since none of the pairs presented all the p -values of the three F-tests significantly lower than 10%.

6.2. Smooth transition model results

Among the 45 pairs of prices analyzed (30 for cross-product and 15 for spatial price ratio), the smooth transition specification was the most appropriate for seven pairs (three for cross-states and four for cross-product price ratio). The final model estimation and their respective coefficients are represented in Tables A1 to A5.

Figs. 3 and 4 illustrate the transition function $G(\cdot)$ overtime and against the transition variable s_t (y_{t-6}). The transition between regimes are fast among every pair as represented by the steep line between

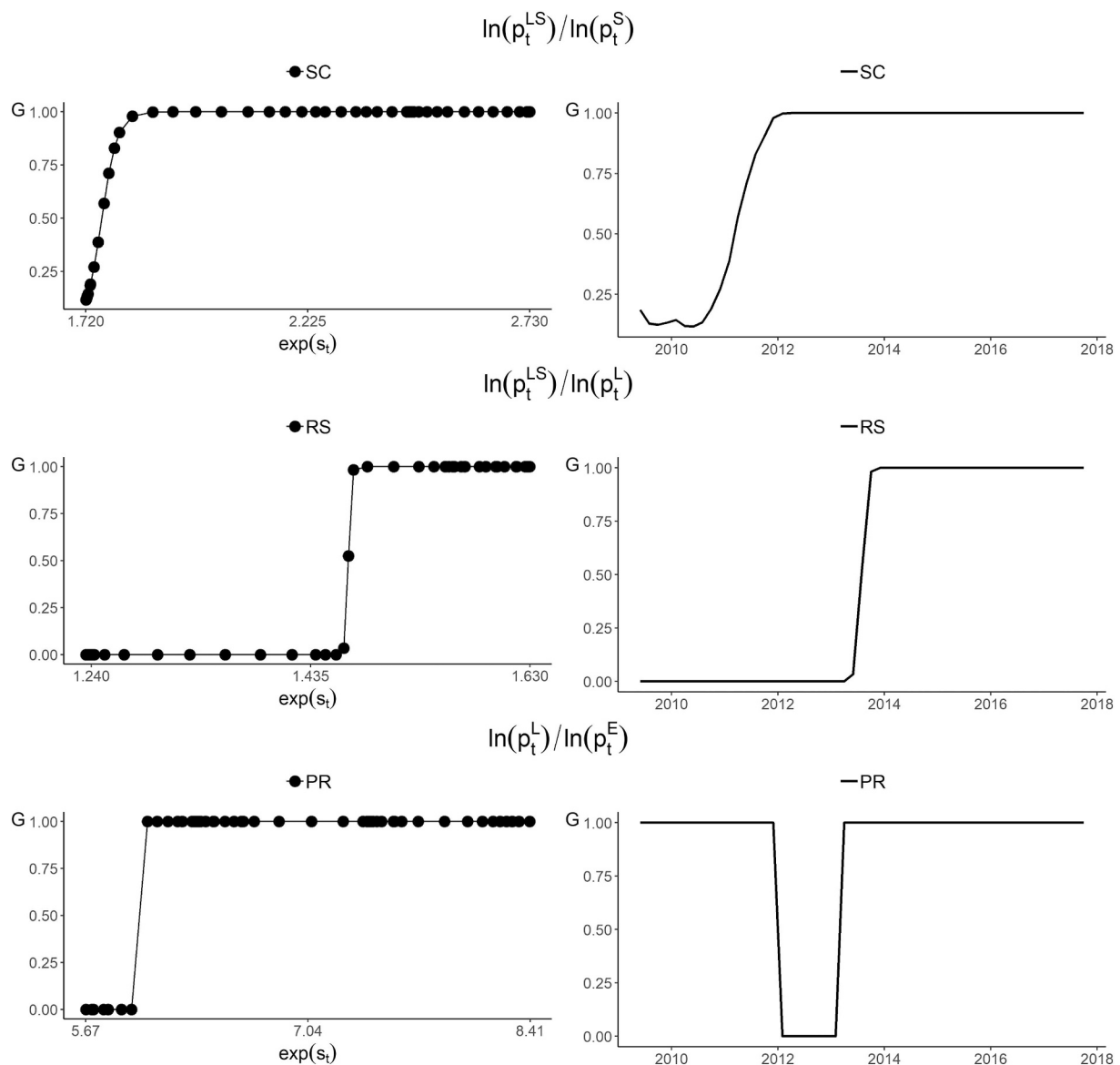


Fig. 4. Spatial Linkage between pine stumpage prices by product. The odd columns show the transition function $G(\cdot)$ versus the respective transition variable (s_t), while the even columns have the transition function $G(\cdot)$ over time (Feb/09 to Jun/17). p_t = stumpage price, PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. FW = Fuelwood, PW = Pulpwood, ST = Sawtimber, V = Veneer and SV = Special Veneer.

$G(\cdot) = 0$ and $G(\cdot) = 1$. The transition to a new regime of either spatial or cross products of special veneer, veneer and sawtimber might be directly linked to the changes in demand of the highest grade of timber. According to the Brazilian Institute of Geography and Statistics (IBGE, 2018), between 2008 and 2017, the demand for sawtimber and higher grades increased 18% (from 13.8 to 16.4 million cubic meters) in PR, 32% (from 7.9 to 10.5 million cubic meters) in SC and 20% (from 5.3 to 6.4 million cubic meters) in RS. Most of this demand comes from the international market, between 2008 and 2017 the exportation of Pine lumber expanded 155% in South of Brazil (from 0.85 to 2.0 million cubic meters) (SECEX Secretariat of Foreign Trade, 2019).

6.3. Generalized impulse response function

The GIRFs of the spatial and cross-product price ratio models are

reported in Figs. 5 and 6 respectively: markets return to equilibrium after an external shock when the GIRF stabilizes. As illustrated there, negative and positive shocks in autoregressive models are symmetric, with fairly fast conversion rates. The return to the equilibrium takes on average 10 months for spatial price pairs, ranging between 6 and 13 months. On the other hand, the autoregressive specification of cross products seems to convert at a slower speed than spatial price linkages. It took on average 12 months to cross-product price ratio returns to equilibrium, with a maximum of 16 months for the price ratio between sawtimber and fuelwood in SC, and a minimum of 8 months for the price ratio between sawtimber and pulpwood in SC.

The linear and non-linear ECM showed a very different outcome. They differentiated from the autoregressive models by two main aspects. First, the GIRFs are not perfectly symmetric. For instance, positive shocks on the price ratio of special veneers in PR and RS have on

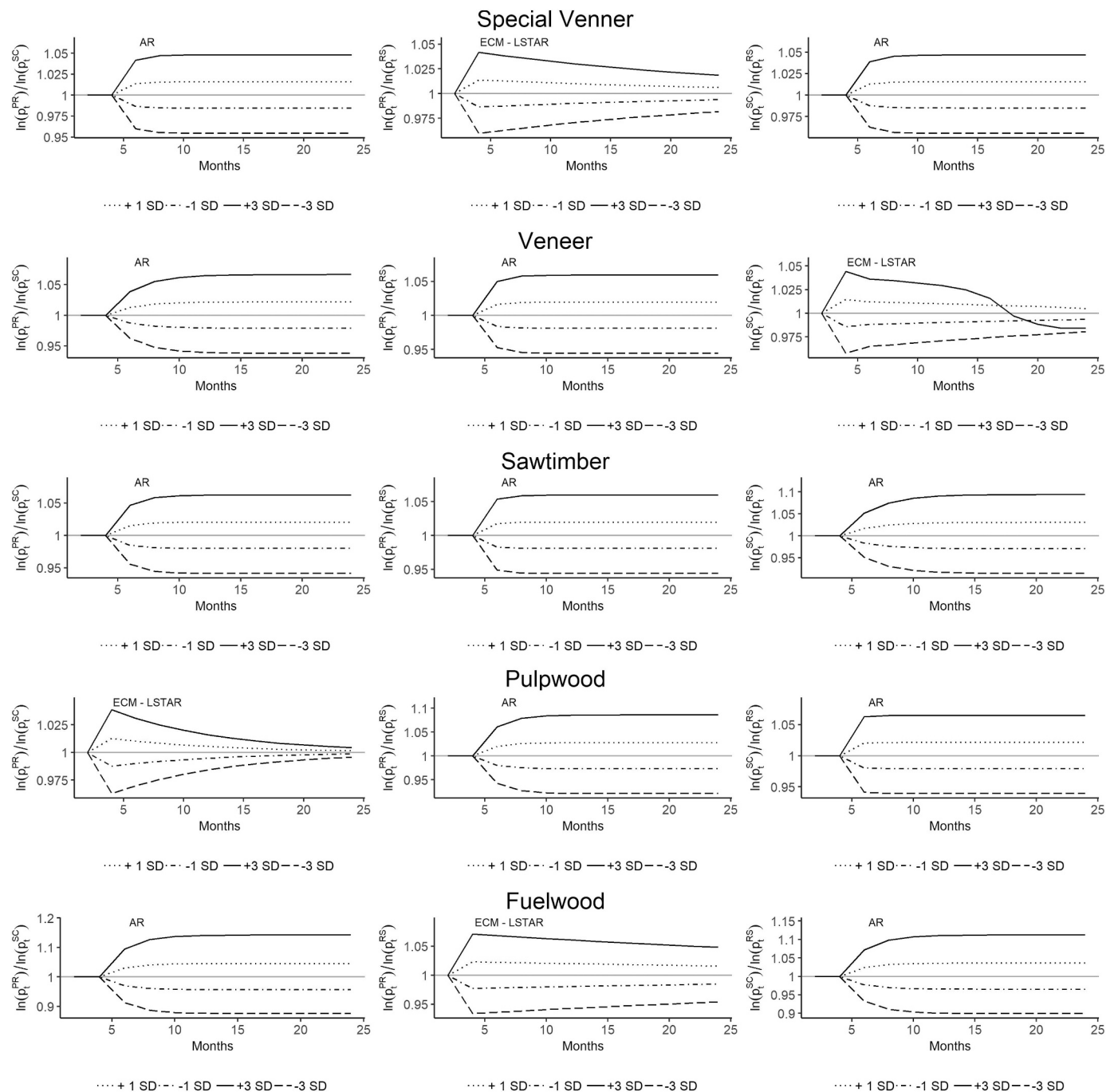


Fig. 5. Generalized Impulse Response Function by Product and spatial pairs. p_t = stumpage price, PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul.

average 4% higher impact than negative shock respectively. Similar results are also found on cross products, a positive shock on the price ratio between veneer and fuelwood in PR has on average 5% higher impact than a negative shock. The second distinction between the ECM and autoregressive models is that the ECM GIRFs tend to not stabilize within the 24 months horizon. The two exceptions are the price ratio between sawtimber and fuelwood, and special veneer and veneer in RS. Both GIRFs stabilized close to 20 months, twofold the conversion speed of autoregressive models.

Another convenient approach to analyze the regime shift on LSTAR and ESTAR models is to impose a threshold to the GIRF. Figs. 7 and 8 break down the GIRF into $G(.) > 0.5$ and $G(.) < 0.5$. All the different price ratio combinations either between states or cross-products show that the speed of convergence is strongly dependent on the regime. For instance, the GIRF of special veneer price ratio between PR and RS converges within 10 months if $G(.) < 0.5$, and longer than 24 months if $G(.) > 0.5$. On the other extreme, external shocks on the price ratio between veneer and fuelwood in the state of PR did not stabilize within

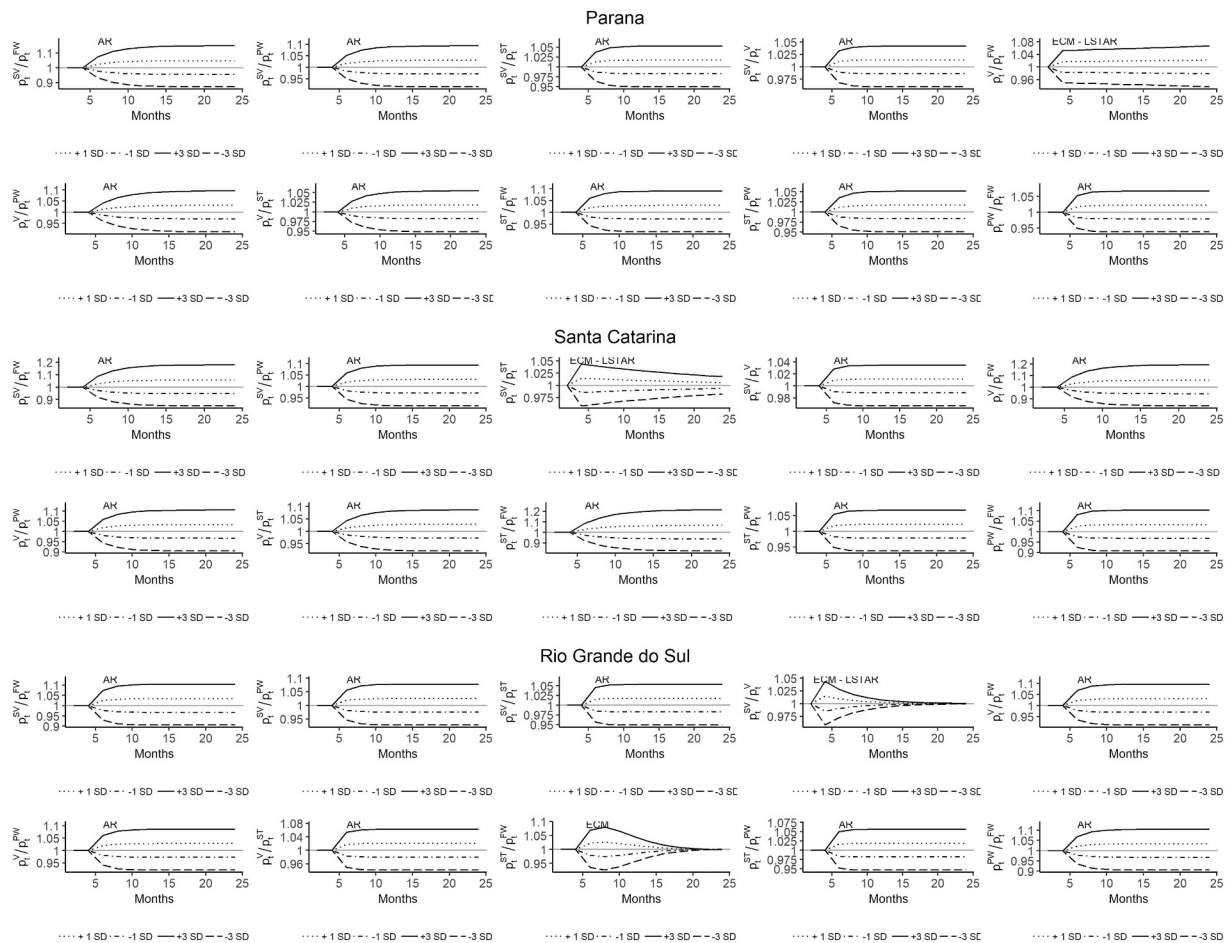


Fig. 6. Generalized Impulse Response Function by Product and spatial pairs. p_t = stumpage price, FW = Fuelwood, PW = Pulpwood, ST = Sawtimber, V = Veneer and SV = Special Veneer.

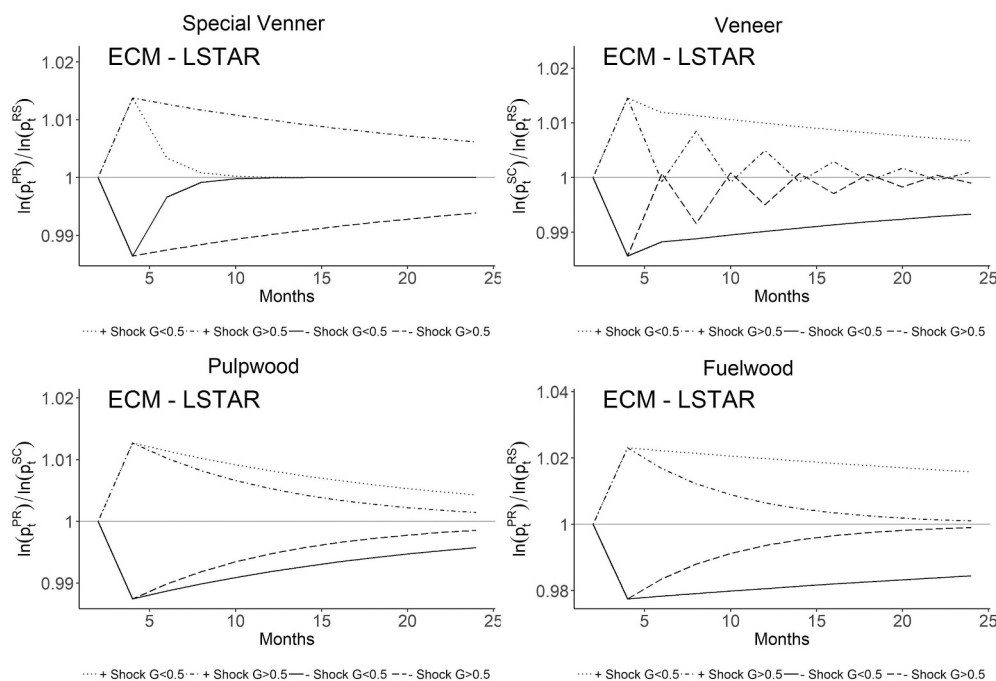


Fig. 7. Conditional Generalized Impulse Response Function by Product and spatial pairs. p_t = stumpage price, PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul.

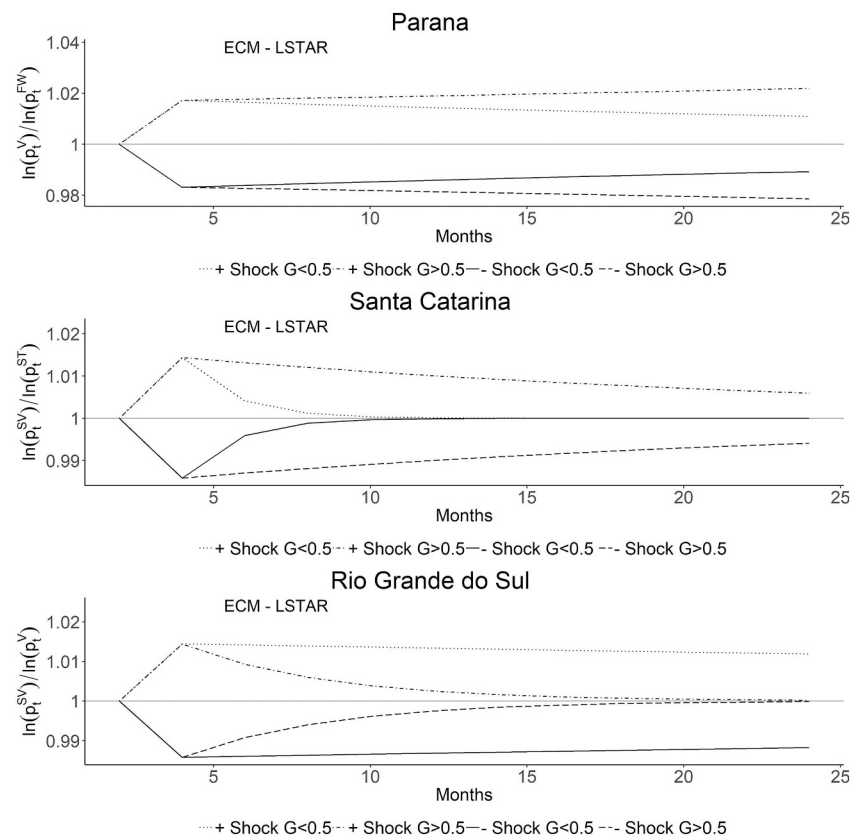


Fig. 8. Conditional Generalized Impulse Response Function by Product and spatial pairs. p_t = stumpage price, FW = Fuelwood, PW = Pulpwood, ST = Sawtimber, V = Veneer and SV = Special Veneer.

24 months when $G(.) < 0.5$, and will not convert at all when $G(.) > 0.5$, implying that under this regime there is no conversion to the equilibrium in the long run.

7. Discussion and conclusion

We investigate the stumpage price linkage for five pine products (fuelwood, pulpwood, sawtimber, veneer, and special veneer) across three states in Brazil. We used a series of tests for linearity and unit root for each pair and evaluated the market integration using a GIRF. Our findings indicate a variety of model specifications with few price pairs presenting a nonlinear behavior. Among other results, the GIRF showed that the linear models converged faster to the equilibrium, indicating a strong market linkage in comparison to the nonlinear counterparts.

Current studies in the United States have also observed cointegration across products (e.g., Størdal and Nyrud, 2003) and strong evidence of spatial cointegration in both markets of pulpwood and sawtimber markets (e.g., Jaunky and Lundmark, 2015; Yin et al., 2002). We also find different levels of market cointegration in the South of Brazil, reinforcing the impact of local supply and demand conditions. In Brazil for instance, the speed of adjustment to a new market equilibrium might be explained by the intensive nature of planted forests in Brazil, where diameter class changes faster than in other countries because of highly productive forests.

The spatial analysis shows comparable results with other studies; our model follows a similar approach like the one proposed by Hood and Dorfman (2015), who studied the cointegration dynamics of sawtimber prices across different regions in the U.S. South using an ECON-

STAR model. Consistent with their findings, we also validated the hypothesis that market regimes vary over time. Their model also shows some agreement with a linear test and, in some cases, the linear model indicated non-cointegration while the STAR model showed partial cointegration and vice-versa.

Although we found some cointegration across states, the level of aggregation in our data (by state) might not reveal how local economic agents truly behave. For instance, tax rate differences across states might prevent trade and therefore reduce market linkages, mainly in low-value products like stumpage. In addition, the costs of harvesting and transporting the timber product across states might be higher than the actual stumpage price, which would make price movements to behave randomly. Chudy et al. (2020) observed similar results in the global roundwood prices. Their results indicate that the majority of roundwood markets across different countries are not linked for several reasons including transaction costs, the different characteristics and end-market uses, transaction costs and legal restriction.

Goodwin et al. (2011) for instance found a strong price linkage between U.S. and Canadian OSB markets. However, stumpage price linkage could increase by using a smaller geographic scale since in-state trade would not face any taxation. Other papers also faced similar limitations. With sawtimber and pulpwood stumpage prices in the U.S. South, Yin et al. (2002) found cointegration between 13 pine sawtimber and 11 pulpwood markets regions, but breaking down to a smaller geographic scale, they recognized, could offer local perspectives. Even though it was not tested, our results did not show any indication of how distance could affect cointegration in the South of Brazil as observed in the US South by Bingham et al. (2003) and Yin et al. (2002).

Our results indicate small differences in cointegration levels between state pairs. Bingham et al. (2003) found that local market capacity plays an important role in determining price linkages. Because of the similar model specifications across different states and products, the GIRF did not show much difference (except between linear and non-linear models). Mei et al. (2010) showed for instance that North Carolina coastal plains, southern and northern Georgia can drive prices in other markets in the U.S. South.

We also observed a different speed of conversion to the equilibrium between cross products and spatial markets. This result is as expected since the temporal relationship between forest products adds uncertainties to future supply and demand; even the well-informed forest managers or investors cannot foresee economic and biological changes within a forest management span. A balanced inventory with multiple products seems to be the safest approach to account for these external shocks. The spatial linkage, on the other hand, is faster observed by market players since changes in nearby markets are easy to notice and to adapt given possible price movements. For instance, a pulp mill inauguration will increase demand for timber, which might change price dynamics in inter-regional markets. However, it is not very clear how the sawlog market would react with the reduction of sawtimber supply in the long run.

Timberland investors in Brazil tend to be more affected by the lack of infrastructure and market inefficiency (often caused by government policies such as cross-state taxes) than in developed countries. Also, Brazil is an attractive market for timberland investors as well as for the pulp and paper industry. Monitoring how prices behave in these countries is essential to reduce risks as well as to understand the supply and demand responses to external shocks. The advanced techniques used here could capture nonlinear adjustments to market equilibrium of

commodities prices.

Our study also has important policy implications. Given the current global protectionism wave, the tax tariff applied between states in Brazil mimics possible consequences of international trade tariffs on timber markets, like arbitrage profits and consequently the misallocation of resources (Johnston and van Kooten, 2017). Policy makers can also use our results to evaluate actions that encourage competition like increasing productivity and reducing bottlenecks in the supply chain. These problems are well known in developing countries and have profound effects on timber prices, and consequently on economic welfare.

Further research should focus on supply and demand quantities instead of only prices, since there is a lack of research on this type of modeling in Brazil and other developing countries. Also, since timberland investors continue to expand their investment to several countries, there is a need to understand if stumpages prices or other market factors can be transmitted to other countries and affect their returns.

Declaration of Competing Interest

None

Acknowledgments

This research was funding by the Conselho Nacional de Desenvolvimento Científico e Tecnológico (CNPq), Brazil, grant ID: 248688/2013-8. We are grateful to STCP Project Engineering Ltda., Brazil, especially Professor Ivan Tomaselli, for providing support during the course of this research.

Appendix

Table A1
STAR model estimates for the bimonthly log-ratio pine stumpage spatial price linkages by product.

	Energy (< 8 cm)			Pulpwood (8–15 cm)			Sawtimber (15–25 cm)		
	$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{PR}/p_t^{SC})$	$\ln(p_t^{RS}/p_t^{SC})$	$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{PR}/p_t^{SC})$	$\ln(p_t^{RS}/p_t^{SC})$	$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{PR}/p_t^{SC})$	$\ln(p_t^{RS}/p_t^{SC})$
Regime 1									
ΔY_{t-1}	0.76*** (0.13)	0.43*** (0.13)	0.38** (0.18)	0.15 (0.13)	0.09 (0.12)	-0.08 (0.16)	0.19* (0.11)	0.36** (0.17)	0.80*** (0.17)
ΔY_{t-2}	-	-	0.38*** (0.14)	-	-	0.28* (0.16)	-	-	-
ΔY_{t-3}	-	-	-0.08 (0.14)	-	-	0.16 (0.18)	-	-	-
Transition									
γ	-1.50*** (0.00)	0.39*** (0.00)	-0.78*** (0.00)	-1.77*** (0.00)	-1.50*** (0.00)	0.48*** (0.00)	-4.83*** (0.00)	0.03*** (0.00)	2.82*** (0.00)
c	-0.07*** (0.00)	-0.05 (0.05)	-0.13*** (0.01)	0.17*** (0.01)	0.19*** (0.00)	-0.05*** (0.01)	0.20*** (0.00)	-0.07*** (0.01)	0.47 (0.59)
Regime 2									
Intercept	-0.04*** (0.01)	-0.15 (0.30)	0.00 (0.00)	0.00 (0.02)	0.07*** (0.02)	0.05*** (0.01)	0.14** (0.06)	0.04*** (0.01)	-0.81 (3.54)
ΔY_{t-1}	-0.03 (0.16)	1.31 (2.32)	0.50*** (0.15)	0.40* (0.22)	0.56** (0.23)	0.18 (0.13)	-0.06 (0.26)	0.12 (0.12)	-32.45 (143.79)
ΔY_{t-2}	-	-	-0.15 (0.17)	-	-	0.17 (0.13)	-	-	-
ΔY_{t-3}	-	-	0.92*** (0.17)	-	-	0.58*** (0.13)	-	-	-
x_{t-1}	-0.41*** (0.13)	-1.83 (3.84)	-0.10* (0.05)	0.00 (0.11)	-0.30*** (0.11)	-0.50*** (0.11)	-0.53** (0.23)	-0.26*** (0.06)	-7.19 (31.51)
RSS	0.03	0.04	0.02	0.02	0.01	0.01	0.01	0.01	0.01
AIC	-200.17	-181.98	-198.90	-220.78	-262.91	-212.40	-234.10	-256.61	-240.50
BIC	-192.39	-174.19	-187.78	-212.99	-255.12	-201.28	-226.31	-248.82	-232.72
ARR	0.78	1.02	0.83	1.02	0.91	0.87	1.03	1.14	0.93

Note: PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. Standard errors are given between parenthesis. RSS denotes the Residual Sum of Squares. AIC is the Akaike Information Criterion and BIC denotes Bayesian Information Criterion. ARR denotes the ratio of the STAR model versus AR model residual standard error. * $p < .1$, ** $p < .05$, *** $p < .01$.

Table A2

STAR model estimates for the bimonthly log-ratio pine stumpage spatial price linkages by product.

	Veneer (25–35 cm)		Special Veneer (> 35 cm)	
	$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{RS}/p_t^{SC})$	$\ln(p_t^{PR}/p_t^{RS})$	$\ln(p_t^{RS}/p_t^{SC})$
Regime 1				
ΔY_{t-1}	-0.01 (0.12)	0.13 (0.11)	0.15 (0.13)	-0.08 (0.16)
ΔY_{t-2}	-	-	-	0.28* (0.16)
ΔY_{t-3}	-	-	-	0.16 (0.18)
Transition				
γ	-1.32*** (0.00)	-1.32*** (0.00)	-1.77*** (0.00)	0.48*** (0.00)
c	0.14*** (0.00)	0.01*** (0.00)	0.17*** (0.01)	-0.05*** (0.01)
Regime 2				
Intercept	0.04 (0.23)	0.01* (0.01)	0.00 (0.02)	0.05*** (0.01)
ΔY_{t-1}	0.39 (2.28)	0.37 (0.30)	0.40* (0.22)	0.18 (0.13)
ΔY_{t-2}	-	-	-	0.17 (0.13)
ΔY_{t-3}	-	-	-	0.58*** (0.13)
x_{t-1}	-0.56 (1.49)	-0.50* (0.24)	-0.15* (0.09)	-0.50*** (0.11)
RSS	0.01	0.01	0.02	0.01
AIC	-268.88	-256.73	-220.78	-212.40
BIC	-261.09	-248.94	-212.99	-201.28
ARR	0.86	0.97	1.02	0.87

Note: PR = Paraná, SC = Santa Catarina, RS = Rio Grande do Sul. Standard errors are given between parenthesis. RSS denotes the Residual Sum of Squares. AIC is the Akaike Information Criterion and BIC denotes Bayesian Information Criterion. ARR denotes the ratio of the STAR model versus AR model residual standard error. * $p < .1$, ** $p < .05$, *** $p < .01$.

Table A3
STAR model estimates for the bimonthly log-ratio pine stumpage cross-product price linkages - Paraná.

	$\ln(p_t^E/p_t^{PW})$	$\ln(p_t^E/p_t^V)$	$\ln(p_t^E/p_t^{SV})$	$\ln(p_t^E/p_t^{ST})$	$\ln(p_t^{PW}/p_t^V)$	$\ln(p_t^{PW}/p_t^{SV})$	$\ln(p_t^{PW}/p_t^{ST})$	$\ln(p_t^V/p_t^{SV})$	$\ln(p_t^V/p_t^{ST})$	$\ln(p_t^{SV}/p_t^{ST})$
Regime 1										
Δy_{t-1}	0.79*** (0.15)	0.74*** (0.09)	0.38** (0.15)	0.58*** (0.13)	10.53 (118.15)	24.36 (191.14)	0.48 (1223.63)	0.04 (0.17)	0.72*** (0.26)	0.56** (0.24)
Δy_{t-2}	-	-	-	-	-	-	-	-	-	-0.16 (0.25)
Δy_{t-3}	-	-	-	-	-	-	-	-	-	-
Transition										
γ	-1.23*** (0.00)	-1.32*** (0.00)	-3.30*** (0.00)	1.65*** (0.00)	0.21*** (0.00)	-1.23*** (0.00)	0.48*** (0.00)	-2.85*** (0.00)	-1.23*** (0.00)	3.54*** (0.00)
c	-0.98*** (0.00)	-1.80*** (0.00)	-2.27 (2.20)	-1.03*** (0.20)	-1.29*** (0.29)	-1.58*** (0.07)	-0.89 (1.47)	-0.35*** (0.00)	0.35*** (0.00)	1.61*** (0.57)
Regime 2										
Intercept										
Δy_{t-1}	-0.05 (0.07)	-0.35** (0.16)	-0.01 (0.10)	-186.05 (752.52)	0.00 (0.01)	-0.01 (0.01)	0.00 (0.01)	0.00 (0.01)	0.00 (0.01)	16.47 (77.36)
Δy_{t-2}	-0.06 (0.13)	-0.63*** (0.21)	0.48*** (0.13)	-1019.61 (4159.80)	0.49*** (0.10)	0.33*** (0.10)	0.28*** (0.10)	0.25* (0.13)	0.41*** (0.11)	-397.49 (1784.52)
Δy_{t-3}	-	-	-	-	-	-	-	-	-	-161.68 (835.93)
x_{t-1}	-0.05 (0.07)	-0.17* (0.09)	0.00 (0.05)	-112.76 (461.06)	0.00 (0.01)	0.00 (0.01)	0.00 (0.01)	0.00 (0.04)	0.00 (0.02)	-15.90 (76.14)
RSS	0.01	0.01	0.03	0.02	0.01	0.01	0.01	0.01	0.00	0.01
AIC	-264.99	-264.96	-224.77	-247.73	-281.99	-260.98	-294.95	-309.91	-325.40	-283.40
BIC	-256.66	-256.63	-216.45	-239.40	-273.67	-252.65	-286.63	-301.58	-317.08	-272.68
ARR	0.94	0.75	0.98	0.96	0.99	0.99	0.99	0.93	0.99	0.97

Note: E = Energy (< 8 cm), PW = Pulpwood (8–15 cm), V = Veneer (15–25 cm), SV = Special Veneer (25–35 cm), ST = Sawtimber (> 35 cm). Standard errors are given between parenthesis. RSS denotes the Residual Sum of Squares. AIC is the Akaike Information Criterion and BIC denotes Bayesian Information Criterion. ARR denotes the ratio of the STAR model versus AR model residual standard error. * $p < .1$, ** $p < .05$, *** $p < .01$.

Table A4
STAR Model Estimates for the Bimonthly Log-Ratio Pine Stumpage Cross-Product Price Linkages - Rio Grande do Sul.

	$\ln(p_t^E/p_t^{PW})$	$\ln(p_t^E/p_t^V)$	$\ln(p_t^E/p_t^{SV})$	$\ln(p_t^E/p_t^{ST})$	$\ln(p_t^{PW}/p_t^V)$	$\ln(p_t^{PW}/p_t^{SV})$	$\ln(p_t^{PW}/p_t^{ST})$	$\ln(p_t^V/p_t^{SV})$	$\ln(p_t^V/p_t^{ST})$	$\ln(p_t^{SV}/p_t^{ST})$
Regime 1										
Δy_{t-1}	140.34 (916.60)	-1.47 (1.39)	0.18* (0.11)	0.35*** (0.12)	0.52*** (0.16)	-0.17 (0.53)	0.37*** (0.13)	-60.16 (434.74)	0.12 (0.10)	0.23* (0.12)
Δy_{t-2}	-	-	-	-	-	-	-	-	-	-
Δy_{t-3}	-	-	-	-	-	-	-	-	-	-
Transition										
γ	1.47*** (0.00)	1.74*** (0.00)	-0.78*** (0.00)	1.56*** (0.00)	-2.58*** (0.00)	-0.06*** (0.00)	1.47*** (0.00)	1.11*** (0.00)	0.30*** (0.00)	1.74*** (0.00)
c	-0.90*** (0.10)	-1.79*** (0.02)	-1.86*** (0.02)	-1.18*** (0.00)	-0.97*** (0.00)	-1.60*** (0.04)	-0.37*** (0.04)	-0.66** (0.27)	0.54*** (0.00)	1.46*** (0.29)
Regime 2										
Intercept										
Δy_{t-1}	-0.04 (0.02)	-0.23*** (0.08)	-25.46 (58.82)	-0.45*** (0.11)	-0.08 (0.08)	-0.01 (0.02)	-82.84 (341.32)	-0.01** (0.01)	0.29 (0.20)	38.09 (116.38)
Δy_{t-2}	0.30** (0.12)	0.42*** (0.10)	-18.36 (58.03)	-0.33 (0.23)	0.08 (0.17)	0.20* (0.11)	16.22 (136.57)	0.51*** (0.12)	-0.51 (0.49)	-43.39 (141.11)
Δy_{t-3}	-	-	-	-	-	-	-	-	-	-
x_{t-1}	-0.05 (0.04)	-0.13*** (0.05)	-12.27 (27.88)	-0.35*** (0.10)	-0.08 (0.09)	0.00 (0.02)	-189.95 (787.42)	-0.03** (0.02)	-0.45 (0.33)	-36.41 (111.26)
RSS	0.03	0.02	0.02	0.02	0.02	0.02	0.01	0.01	0.02	0.01
AIC	-232.42	-237.22	-234.97	-238.80	-244.35	-251.43	-268.30	-270.50	-258.27	-279.15
BIC	-224.09	-228.89	-226.65	-230.47	-236.02	-243.10	-259.98	-262.18	-249.94	-270.83
ARR	0.97	1.01	0.87	0.89	0.95	0.97	0.91	0.95	0.95	0.94

Note: E = Energy (< 8 cm), PW = Pulpwood (8–15 cm), V = Veneer (15–25 cm), SV = Special Veneer (25–35 cm), ST = Sawtimber (> 35 cm). Standard errors are given between parenthesis. RSS denotes the Residual Sum of Squares. AIC is the Akaike Information Criterion and BIC denotes Bayesian Information Criterion. ARR denotes the ratio of the STAR model versus AR model residual standard error. * $p < .1$, ** $p < .05$, *** $p < .01$.

Table A5
STAR model estimates for the bimonthly log-ratio pine stumpage cross-product price linkages - Santa Catarina.

	$\ln(p_t^E/p_t^{PW})$	$\ln(p_t^E/p_t^V)$	$\ln(p_t^E/p_t^{SV})$	$\ln(p_t^E/p_t^{ST})$	$\ln(p_t^{PW}/p_t^V)$	$\ln(p_t^{PW}/p_t^{SV})$	$\ln(p_t^{PW}/p_t^{ST})$	$\ln(p_t^V/p_t^{SV})$	$\ln(p_t^V/p_t^{ST})$	$\ln(p_t^{SV}/p_t^{ST})$
Regime 1										
ΔY_{t-1}	143.59 (662.84)	0.42*** (0.13)	0.40*** (0.10)	-0.53 (4.21)	4.91 (41.13)	-10.75 (67.41)	0.44** (0.21)	0.03 (0.11)	0.80*** (0.26)	1.15*** (0.33)
ΔY_{t-2}	-	0.09 (0.31)	-	1.07 (4.75)	-	-	0.09 (0.20)	-	-1.04*** (0.40)	-0.59 (0.40)
ΔY_{t-3}	-	0.08 (0.31)	-	1.43 (3.53)	-	-	-0.21 (0.20)	-	-	-
Transition										
γ	0.66*** (0.00)	-5.55*** (0.00)	-0.51*** (0.00)	3.54*** (0.00)	-0.60*** (0.00)	3.81*** (0.00)	-2.85*** (0.00)	-1.05*** (0.00)	2.82*** (0.00)	-0.51*** (0.00)
c	-0.97*** (0.12)	-1.70*** (0.00)	-1.88*** (0.04)	-1.29 (5.95)	-1.34*** (0.14)	-8.48 (11.64)	-0.65*** (0.00)	-0.37*** (0.00)	0.84 (0.94)	0.58*** (0.01)
Regime 2										
Intercept										
ΔY_{t-1}	-0.03* (0.02)	-0.14** (0.05)	-0.02 (0.34)	-0.12 (0.38)	-0.01 (0.01)	-0.01 (0.02)	-0.10*** (0.02)	0.00 (0.04)	0.38 (1.74)	0.09*** (0.02)
ΔY_{t-2}	-0.04 (0.16)	1.08*** (0.17)	0.60 (0.45)	1.72 (3.30)	0.37*** (0.10)	0.57 (1.40)	0.20** (0.10)	0.90** (0.38)	-2.84 (14.83)	0.11 (0.12)
ΔY_{t-3}	-	-0.56*** (0.15)	-	-1.68 (4.43)	-	-	0.19* (0.11)	-	9.46 (42.55)	-0.32*** (0.10)
x_{t-1}	-	0.42*** (0.12)	-	-0.52 (3.23)	-	-	0.67*** (0.11)	-	-	-
RSS	-0.05*** (0.02)	-0.09*** (0.03)	0.00 (0.19)	-0.09 (0.30)	0.00 (0.01)	0.00 (0.01)	-0.17*** (0.04)	0.00 (0.11)	-0.76 (3.46)	-0.09*** (0.02)
AIC	0.03	0.03	0.04	0.04	0.02	0.02	0.01	0.00	0.01	0.01
BIC	-217.50	-194.65	-209.90	-192.03	-246.81	-249.24	-245.92	-323.27	-272.79	-272.23
ARR	-209.17	-182.24	-201.57	-179.62	-238.49	-240.91	-233.51	-314.94	-262.07	-261.51
	0.94	0.86	0.96	0.91	0.99	0.99	0.85	0.92	0.91	0.85

Note: E = Energy (< 8 cm), PW = Pulpwood (8–15 cm), V = Veneer (15–25 cm), SV = Special Veneer (25–35 cm), ST = Sawtimber (> 35 cm). Standard errors are given between parenthesis. RSS denotes the Residual Sum of Squares. AIC is the Akaike Information Criterion and BIC denotes Bayesian Information Criterion. ARR denotes the ratio of the STAR model versus AR model residual standard error. *p < .1, **p < .05, ***p < .01.

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