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Substitutability of U.S. and Canadian softwood lumber: A threshold modeling approach

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ABSTRACT

The U.S.-Canada softwood lumber dispute stands as the most protracted trade conflict between the two countries, with its history dating back to the early 1980s. U.S. lumber producers claim that Canadian lumber imports are unfairly subsidized, thereby, harming the domestic lumber industry. The U.S. has imposed countervailing and antidumping duties on Canadian softwood lumber imports to protect the domestic market. An important aspect of this trade dispute is the substitutability of lumber products between the two countries. Canadian lumber imports are detrimental to the U.S. industry only when the products are substitutes. In such a scenario, the imposition of trade restrictions would have a significant influence on the domestic market. Motivated by this consideration, we investigate dynamic relationships between prices for Spruce-Pine-Fir, a softwood species primarily imported from Canada, and Southern Yellow Pine, one of the most popular softwood species domestically produced in the U.S. We apply threshold models in an evaluation of the degree of substitutability. Our empirical findings indicate that within the $+/-$ 3.4 % price differential band, the two lumber products demonstrate a higher degree of substitutability.

1. Introduction

The United States (U.S.) holds a prominent position as one of the world's largest producers and consumers of softwood lumber. According to data compiled by the U.S. International Trade Commission (USITC), softwood lumber consumption in 2022 reached nearly 53 billion board-feet, a 12.8 % increase from the 2016 consumption of 47 billion board-feet (USITC, 2023). A significant share of the softwood lumber consumed in the U.S. is sourced from Canada. The USITC reports that in 2022, U.S. producers had a market share of 68.7 %, while Canada's share was approximately 24.1 %, leaving the remaining 7.2 % for other exporting markets. Additionally, in 2015, approximately 41 % of U.S. lumber production and 95 % of Canadian production came from species that potentially have common end-use applications in both countries (USITC, 2023).

Considering the significant contribution of Canadian exports to U.S. softwood lumber consumption and the possible overlapping applications in lumber species, the objective of this study is to address the level of substitutability between softwood lumber products in the U.S. and Canada. Substitutability between imports and domestic-like products is a crucial aspect of international trade analysis, as it determines the efficacy of trade-restricting policies on both imports and domestic markets (Gan, 2006). If softwood lumber products from Canada are considered substitutes for domestically produced lumber in the U.S., the influx of foreign products with lower prices into the domestic market could harm the U.S. industry. In such a scenario, the imposition of import restrictions would become a desirable protection measure for U.S. producers, albeit with detrimental effects on U.S. consumers.

There has been a longstanding trade issue between the U.S. and Canada in the softwood lumber industry, known as the U.S.-Canada Softwood Lumber Dispute. Over the history of the dispute, the U.S. has imposed several trade restrictions on Canadian softwood lumber imports, including antidumping duties (AD) and countervailing duties (CVD).[1] Although various short-term agreements, typically effective for five to nine years, have been implemented to temporarily alleviate the conflict, there was no lasting resolution established, leaving the ongoing

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[1] The term "dumping" denotes the act of selling lumber to a foreign country at a price below the production cost or selling at a price lower than domestic prices. Allegations of dumping were raised by U.S. lumber producers, contending that the price of Canadian imports to the U.S. market is lower than the production cost in Canada.

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friction between the two countries unresolved. The persistence of tension in the bilateral lumber trade can be attributed, in part, to inherent disparities in population levels, and the composition of the lumber market. The U.S. and Canada possess abundant forest resources, but their distinct population sizes and development pressures result in markedly different lumber market scales (Hoover and Fergusson, 2018). As the U.S. stands out as the largest producer, importer, and user of lumber, and Canada being its primary exporter, these two countries are particularly susceptible to any changes in trade policies and exogenous shocks from the other market (Devadoss, 2009).

This dispute is centered around two issues. The first issue is grounded in the fundamental differences in forestry management practices between the two countries, for which subsidies received by the Canadian lumber industry are a central point of contention (Rahman and Devadoss, 2002). Approximately 93.3 % of forestland in Canada is publicly owned (Canada Forest Service, 2022), compared to 42 % of federally and state-owned forestland in the U.S. (Hoover and Riddle, 2021).[2] Since the majority of the timber is harvested from provincial crown land in Canada, U.S. lumber producers claim that Canadian stumpage fees, the prices landowners charge for the right to harvest timber, are not determined in a competitive market. This, according to their argument, results in artificially low prices for timber produced in Canada (Zhang, 2007, Ch 3). In contrast, U.S. timber is typically harvested from privately owned lands, where stumpage fees are determined through competitive auctions (Sedjo, 2006). Subsidizing an exported commodity may violate the World Trade Organization (WTO) Subsidies Agreement, thereby providing U.S. producers with a case against Canadian exporters within the WTO as justification for trade restrictions against the softwood lumber product category.

The second issue fueling the dispute is the substitutability of imported and domestic lumber. U.S. producers argue that Canadian lumber imports are good if not perfect substitutes for domestically produced lumber products. This means that domestic and imported lumber in the U.S. directly competes with one another. To the extent that subsidies underlie Canadian production, this results in unfair competition thereby with domestic producers and harm to the U.S. lumber industry. The issue of substitutability is an important competition factor in the threat of injury analysis (USITC, 2019). According to provisions outlined in the U.S. code §1671, if the administering authority concludes that a countervailable subsidy–whether directly or indirectly–has been provided to the production of a commodity imported to the U.S. and such subsidization is causing material injury or threatening material injury to a U.S. industry, a CVD may be imposed on the concerned merchandise. The amount of the CVD is determined to be equal to the net countervailable subsidy (19 USC. § 1671(a)). The imposition of an AD requires the establishment of two similar determinations. The administering authority must determine that foreign merchandise is being sold in the U.S. at a rate below its fair market value, and a U.S. industry is experiencing material injury or is under the threat of material injury (19 USC. § 1673).

Despite contentions to the contrary raised by U.S. producers, the Canadian government responds to the allegation by stating that lumber is used in a wide range of industries and, therefore, the policy does not qualify as a subsidy under the U.S. law (Rahman and Devadoss, 2002). Moreover, U.S. consumers claim that domestically produced lumber and Canadian lumber are not substitutes because the types of lumber used for housing framing serve different purposes and are not interchangeable (Nagubadi et al., 2004). These perspectives assert that the trade restrictions imposed on Canadian lumber imports are motivated more by protectionist sentiments than by an objective response to perceived unfair practices (*Montreal Economic Institute*, 2022). According to economists at the Montreal Economic Institute (MEI), the placement of tariffs harms the economies of both countries, and such protectionist measures act as barriers to free trade (Rancourt and Giguère, 2022). Their analysis indicates that the financial impact of trade restriction policies during the softwood lumber dispute is 26 times more significant for the U.S. population compared to Canadian producers. This results in an estimated $1.5 billion reduction in the economic welfare of U.S. consumers (Rancourt and Giguère, 2022). Furthermore, the decrease in welfare for U.S. consumers is projected to exceed the gains experienced by U.S. producers by 7 % (Rancourt and Giguère, 2022).

Various studies have delved into assessing the welfare impacts arising from trade conflicts within the forestry sector, with scholars finding mixed support for trade regulatory policies based on statistical estimations. For instance, Buongiorno and Johnston (2018) found that import barriers on foreign lumber increase the welfare of U.S. producers, but the increase was less than the losses incurred by U.S. consumers. On a similar note, Johnston and van Kooten (2017) estimated the effects of the tariff rate quota used in the 2006 Softwood Lumber Agreement (SLA2006), which resulted in an annual deadweight loss of $28 million, borne by U.S. consumers. Johnston and van Kooten (2017) further noted that the absence of SLA2006 would increase the supply share of Alberta while improving aggregate welfare. van Kooten and Johnston (2014) observed that there would be a very small change in U.S. welfare from the removal of export restrictions on Canadian lumber, estimated at around $16 million, while Canada would gain $91.8 million. Results from Johnston and Parajuli (2017) indicated that the imposition of a 26.75 % CVD and AD tariff on Canadian lumber imports would increase U.S. lumber production by 2.6 million m^3 with a price increase of $5–7/$m^3$, while the Canadian lumber exports to the U.S. fall by 4.9 million m^3.

The formulation of trade policies is complex and often requires balancing the interests of the executive and legislative branches. Confirming subsidies on Canadian lumber is beyond the scope of this study. In addition, instead of estimating the welfare changes associated with existing trade duties, this study focuses on examining the substitutability between lumber products from the perspective of Canadian and U.S. lumber price parity. Specifically, our focus is on understanding how price transmission between U.S. and Canadian softwood lumber products varies when price differentials are narrow versus wide. Theoretically, if the two products are close substitutes, we would expect rapid adjustments in market prices in response to deviations in price differentials.

Earlier studies on substitutability between lumber species in North America yield varying results, primarily focusing on calculations of price elasticities of substitution. As noted in the USITC's latest report on investigations of Canadian softwood lumber products, the estimated substitution elasticity between U.S.-produced softwood lumber and subject-imported softwood lumber from Canada ranges between 2 and 5 depending on species and purchasers' preferences (USITC, 2023). Based on the evidence, the Commission considers the domestic product and subject imports from Canada to have "at least moderate degree of substitutability" (USITC, 2023). Nagubadi et al. (2004) investigated the relationship among six lumber species groups in the U.S., including ones imported from Canada, using a restricted translog subcost function. They computed Allen partial elasticities of substitution and found that the Spruce-Pine-Fir group from Canada is a substitute for untreated Southern Yellow Pine and structural panels. Gan (2006) calculated the Armington elasticity of substitution between U.S. domestic and imported forest products, finding that elasticity estimates are generally small and stable, although the aggregated level estimates are increasing, which indicates increasing substitutability of imports for domestic forest products over time. Shook et al. (2007) estimated the cross-price elasticity of demand for U.S. softwood lumber species and did not find strong

[2] The total forestland in Canada is approximately 894 million acres, with about 75.4 % of it being provincially owned, 13.0 % territorial, 2.1 % indigenous, and 1.7 % federal, leaving 6.7 % privately owned land (Canada Forest Service, 2022). In contrast, the total forestland in the U.S. is approximately 765 million acres. As of 2017, 31 % of U.S. forests are federally owned, with state and local governments owning 11 %. In addition, 38 % is owned by private noncorporate households, while private corporations own 20 % (Hoover and Riddle, 2021).

evidence that Canadian softwood species behave as substitutes for U.S. domestic species. Roman et al. (2006) employed a linear approximate almost ideal demand system for U.S. softwood lumber imports from various countries. They found that compared to Canadian lumber imports, the elasticities of substitution from other countries are higher, implying a higher degree of substitutability with U.S. domestic products.

This study contributes to the literature by applying threshold models and examining the dynamic price behaviors between U.S. and Canadian lumber products, filling a gap in the relatively limited research on lumber substitutability. While threshold models are widely used to assess market linkages across different dimensions—such as spatial, market form, and temporal linkages—this study applies these models to examine price transmission between two similar products at the same stage of the supply chain, with shared end-use characteristics.

Traditionally, price transmission is analyzed within the context of market integration, where external shocks in one market can influence prices in another (von Cramon-Taubadel and Goodwin, 2021). Market integration for a commodity can be examined both vertically and horizontally. Vertical market integration examines price relationships at different levels of the supply chain, while horizontal integration assesses the extent of market deviation across different locations for identical products (for example, see Goodwin and Holt, 1999; Goodwin and Piggott, 2001; Bekkerman et al., 2013; Goodwin et al., 2017; Ramsey et al., 2021; Goodwin et al., 2021). In general, arbitrage arises when the price difference between two markets——whether spatially or vertically integrated——exceeds the arbitrage costs (e.g., transaction costs). In response, rational agents will buy in the lower-priced market and sell in the higher-priced market, thereby reducing the price differential and restoring market equilibrium. This process establishes a"neutral band" (delineated by a threshold) within which price differentials do not exceed the costs of arbitrage, resulting in prices within this band being less sensitive to shocks from the other market (Goodwin, 2024).

In the context of substitutability, when deviations from price parity exceed the identified thresholds, we may observe a shift in consumers' consumption patterns rather than typical arbitrage behavior that transfers products between markets. Our empirical analysis focuses on Southern Yellow Pine (SYP), a leading softwood species produced in the U.S., and Spruce, Pine, and Fir (SPF), the predominant softwood species imported from Canada.[3] Given that both lumber species groups take major shares of U.S. consumption, their fungibility remains a focal point in discussions concerning trade policies for Canadian lumber. SPF has a high strength-to-weight ratio, meaning the lumber is lighter but can hold a large amount of weight (AIFP Lumber, 2019, 2020). It also has superior nail-holding capacity and is less susceptible to warping compared to SYP. These characteristics of SPF make it preferred by builders for wall framing. Conversely, SYP's denser fiber composition makes it heavier and more suitable for structural applications such as trusses, beams, and joists (Nagubadi et al., 2004). These distinctive characteristics further complicate the substitutability dynamics between these species as price differentials widen.

Existing studies on market integration within the U.S. and Canadian softwood market restrict their attention to evaluating the effectiveness of the lumber trade dispute. For instance, Baek and Yin (2006) combined a structural simultaneous system of equations with a nonstructural vector autoregression model to assess the impacts of the Memorandum of Understanding (MOU) and Softwood Lumber Agreement (SLA). Their findings suggest that while the MOU effectively regulated lumber imported from Canada, the SLA had no significant impact on restricting these imports or increasing U.S. production. Nagubadi and Zhang (2013) investigated the determinants of U.S. softwood lumber imports from Canada, using a cointegration approach to examine the effects of various trade restriction measures. Their results indicate that all trade restriction measures had imposed a significant negative impact on U.S. imports from Canada. Further, Guo and Johnston (2021) examined the spatial price transmission between Canadian and U.S. markets, comparing the speed of price adjustment across free-trade and trade-restricting periods. They found that trade liberalization reduces domestic-to-foreign transaction costs while increasing domestic transaction costs.

The novelty of our research lies in its focus on market integration to provide new insights into the substitutability of lumber products, which also adds to our understanding of how markets are integrated in the context of cross-commodity price transmission. While our estimation strategy aligns with traditional market integration analyses, we apply threshold models with a distinct motivation. We hypothesize that the substitutability of U.S. and Canadian softwood lumber products may depend on the differences in prices for the alternative products. When prices are comparable, consumers have the flexibility to select the product that best aligns with their individual preferences. In this scenario, the products become more substitutable. However, when price differences are substantial, the consumption incentive shifts toward the lower-cost product as the price benefits dominate, and the ability to substitute is diminished. The varying perceptions of interchangeability between the two products are supported by evidence from the USITC's surveys and questionnaires, which reveal that builders, as reported by the National Association of Home Builders, view them as "sometimes interchangeable" (USITC, 2023). While the questionnaire does not specify the reasons for these perceptions, price differentials likely play a significant role.

To address the potential presence of threshold effects in the context of lumber substitutability, we employ a Self-Exciting Threshold Autoregressive (SETAR) model, taking the absolute value of the logged price differentials as a forcing variable—that is, the variable whose values cause shifts across alternative substitution regimes. Although cointegration was not the focus of our analysis, we consider bivariate Threshold Vector Error Correction (TVEC) models for the paired-price system to jointly evaluate the interactions among prices, providing a more comprehensive assessment of price dynamics.[4]

The primary findings of our analysis can be summarized as follows. The SETAR model confirms the presence of threshold effects, with an estimated threshold band of $+/-0.034$, or $+/-3.4$ %. Both standard and generalized impulse response analyses reveal a significantly faster market response to shocks within the threshold band, suggesting a higher degree of market integration between the two products when price differentials are narrow. This supports our hypothesis that when price differentials between the two lumber species fall outside the threshold band, the markets operate more independently, and substitution behaviors are more likely to occur within the observed threshold where prices are closer to one another. These insights contribute to our understanding of price transmission dynamics between the U.S. and Canadian softwood lumber markets, providing valuable implications for policymakers, particularly in the context of trade restrictions related to the ongoing softwood lumber dispute.

2. History of the softwood Lumber dispute

The modern history of the U.S.-Canada softwood lumber dispute can

[3] We acknowledge that Douglas Fir, a softwood species from the U.S. Pacific Northwest, is a potential substitute for Canadian SPF. However, its role in the traded lumber market is relatively modest. According to the latest data from the USDA Forest Service's Forest Inventory and Analysis Timber Product Output survey, Douglas Fir accounts for approximately 12,224 million board feet of production, representing around 10.5 % of total U.S. softwood production. In contrast, SYP dominates with approximately 83,962 million board feet, accounting for about 72.3 % of total U.S. softwood production. Therefore, we consider SYP prices to be much more representative and relevant for analyzing market integration and price dynamics in the U.S. softwood lumber market.

[4] See Sun and Ning (2014) and Guo and Johnston (2021) for assessing spatial price linkage among softwood lumber markets in North America.

be divided into five phases: Lumber I (1981–1983), Lumber II (1986–1991), Lumber III (1991–1994), Lumber IV (2001–2006), and Lumber V (2015-present). Among these phases, three short-term settlement agreements were signed to address the dispute, yet a lasting and comprehensive resolution has remained unattainable throughout the historical evolution of the conflict up to the present day.

In 1981, a group of U.S. lumber producers petitioned the U.S. Department of Commerce (USDoC) and the U.S. International Trade Commission (USITC) for an investigation into Canadian lumber imports to the U.S. for a possible countervailing duty (CVD) (Gorte and Grimmett, 2006). This first investigation is known as Lumber I. These producers claimed that Canadian softwood lumber products were unfairly subsidized by the government, allowing Canadian lumber producers to sell at a price significantly below the market value of timber (Rahman and Devadoss, 2002). While the USITC initially collected evidence indicating that Canadian lumber imports were causing harm to the U.S. lumber industry, in 1983, the International Trade Administration (ITA) of the USDoC determined that the subsidies amounted to *"de minimis"* (less than 0.5 %) and that Canadian stumpage programs were not specific to a single industry, making them non-countervailable (Lindsey et al., 2000). As a result of this negative CVD determination, the initial CVD investigation concluded without imposing any duties on Canadian softwood lumber imports (Zhang, 2007).

After Lumber I, the U.S. lumber industry group, the Coalition for Fair Lumber Imports (Coalition), initiated another countervailing duty (CVD) investigation into Canadian softwood lumber imports in 1986. This time, the ITA reversed its previous determination and found evidence that Canadian lumber producers had received a subsidy of 15 % ad valorem (Gorte and Grimmett, 2006; Zhang, 2007).[5] One day before the final CVD determination, the two countries negotiated and signed a Memorandum of Understanding (MOU). Under the terms of the MOU, the Canadian government committed to imposing a 15 % tax on lumber exported to the U.S. for the subsequent five years. The Coalition withdrew its petition immediately after the signing of the MOU (Rahman and Devadoss, 2002). This marked the first legal agreement of the ongoing dispute.

The third investigation of Canadian softwood lumber imports, or Lumber III, began with the termination of the MOU when the Canadian government withdrew in September 1991. In the following October, the U.S. Trade Representative (USTR) self-initiated an investigation upon the pending completion of the previous CVD investigation. This investigation confirmed substitutability between lumber species groups in the U.S. market and concluded with a final determination that Canadian softwood lumber causes material injury to the U.S. industry, and exports are countervailable. As a result, the U.S. government imposed a 6.5 % ad valorem duty on Canadian lumber imports (Congressional Research Service, 2001). Canada appealed the final determination of material injury and the final CVD of 6.5 % under the U.S.-Canada Free Trade Agreement (FTA) (Zhang, 2007). This phase of the dispute was settled with a final FTA panel decision ordering the DoC to revoke the CVD order (Gagné, 2003).

In 1996, both countries were interested in finding a durable solution to the dispute. Upon numerous negotiations, the first softwood lumber agreement emerged, known as the 1996 Softwood Lumber Agreement (SLA1996). Under SLA1996, a tariff-rate quota of 14.7 billion board-feet was established for Canadian softwood lumber imports produced in British Columbia, Alberta, Ontario, and Quebec (Gagné, 2003; Zhang, 2007). The Canadian government was responsible for collecting fees on over-quota shipments, in return, the U.S. government would alleviate any trade restrictions throughout the agreement. However, when SLA1996 expired in March 2001, the countries failed to reach a new agreement, leading to the fourth phase of the dispute.

The Coalition submitted another petition for both CVD and Anti-dumping Duty (AD) shortly after the expiration of SLA1996. In August 2001, the DoC announced an initial CVD of 19.3 % and an AD of 3.1 % on Canadian softwood lumber imports, and the final determination of CVD and AD amounted to 27.2 % in May 2002 (Yin and Baek, 2004; Zhang, 2007).[6] In April 2006, the U.S. and Canada reached a framework agreement to form an end to the dispute. Under this tentative agreement, Canadian lumber products would account for no more than 34 % of the U.S. lumber market share, and an export tax would be imposed on lumber products exported to the U.S. with a price lower than $355 per thousand board-feet (CBC News, 2006). In addition, the U.S. agreed to refund 80 % of the $5 billion duties collected in the previous four years and to withhold from initiating trade restrictions or imposing new tariffs (CBC News, 2006). This agreement officially took effect on October 12, 2006, and was known as the 2006 Softwood Lumber Agreement (SLA2006). It was to last seven years, with a possible two-year extension.

As of the writing of this paper, no further agreement has been reached since the expiration of SLA2006 on October 12, 2015. From 2016 to 2023, the trade dispute has escalated on several occasions with various duty rates imposed over the years. In July 2023, The DoC set the rate on Canadian softwood lumber imports at 7.99 % (Global Affairs Canada, 2023). The Canadian government appealed under dispute-resolution provisions in the North American Free Trade Agreement (NAFTA) to the World Trade Organization (WTO) to challenge the duties. Most recently, the USITC issued its five-year (sunset) review decision on the newest investigation on Canadian softwood lumber on November 30, 2023. All four commissioners arrived at the conclusion that lumber imports from Canada are unfairly subsidized, therefore, remaining a threat to the U.S. lumber industry. Their decision supports the continuation of the countervailing and anti-dumping duty on Canadian softwood imports (Johnston and Parajuli, 2017; USITC Office of External Relations, 2023). In light of the ongoing trade restrictions and continued disputes, our research is particularly relevant, offering critical insights to support policymakers in navigating this complex and unresolved issue.

3. Econometric procedure

In conventional market integration analysis, arbitrage typically refers to the reallocation of trade between spatially distinct markets (Goodwin et al., 2021; Goodwin, 2024). However, in the context of substitutability, market integration is characterized by shifts in consumption from one product to another instead of spatial shifts in trade. Considering two price series of lumber produced in Canada and the U.S., p_t^1 and p_t^2. The linear regression procedure among prices models the long-run relationship is given by

$$p_t^1 = \alpha + \beta p_t^2 + u_t \tag{1}$$

where the residual u_t is a white-noise error term. Note that if both prices are in logarithmic terms, the coefficient β represents the price transmission elasticity ($\partial p_t^1 / \partial p_t^2$), which indicates the responsiveness or sensitivity of one price to changes in another. A price transmission elasticity of one indicates perfect market integration, whereby shocks in one market are instantaneously and fully transmitted to the other market.

The parity relationship between two lumber prices can be described using an autoregressive process:

$$\widetilde{p}_t = \delta \widetilde{p}_{t-1} + \nu_t \tag{2}$$

[5] The reversal of determination was based on a new guideline issued by the DoC, outlining which subsidies are subject to the CVD laws (Gorte and Grimmett, 2006).

[6] The Commerce made a final determination of a 18.79 % CVD and 8.43 % AD on Canadian softwood lumber imports (Zhang, 2007, Ch 8).

where $0 < \delta < 1$ and $\widetilde{p}_t = |p_t^1 - p_t^2|$ is defined as the forcing variable, which triggers a regime-switching behavior when it passes the threshold value, and ν_t is a white-noise error term. We focus on the absolute value of price differentials to examine how significant deviations from a threshold, regardless of direction, influence market behaviors and trigger regime changes. This specification enforces a symmetric threshold for price differentials.

As mentioned earlier, because of consumer preferences for types of lumber products purchased, consumers' responses to price disparities ($\widetilde{p}_t$) may not be linear. To test for the presence of nonlinear and asymmetric behaviors in $\widetilde{p}_t$, we implement four linearity tests, each testing the null hypothesis that the time series follows a linear autoregressive (AR) process against threshold alternatives. The first test is a neural network test for neglected nonlinearity by Lee et al. (1993). This neural network test uses a single hidden layer network, and the null hypothesis of linearity is not rejected if the optimal network weights are zero. The second test is the Keenan (1985) One-Degree Test for nonlinearity, which is based on the correlation of estimated residuals and the squared forecast. If the squared forecast has additional forecasting ability, i.e., correlated with the residuals, then the process is considered nonlinear. The third test, introduced by Tsay (1986), is similar to Keenan's test as it directly tests departures from linearity in the mean. Lastly, we perform the likelihood ratio test for threshold nonlinearity introduced by Chan (1993). If the null hypothesis is not rejected, indicating linearity, the price differentials should be modeled using a linear AR model. Otherwise, threshold models are implemented if the null hypothesis is rejected, suggesting nonlinearity.

With the confirmation of nonlinearity, we evaluate the degree of price transmission by applying a Self-Exciting Threshold Autoregressive (SETAR) model to the absolute value of the price differentials. The SETAR model, developed by Tong (1983, 1990) and Chan (1993), extends the linear AR framework by allowing the autoregressive coefficients to vary based on a data-driven threshold, denoted by θ. Unlike a standard Threshold Autoregressive (TAR) model, which relies on an external variable as the forcing variable, a SETAR model uses the lagged value of the series itself as the forcing variable, making the regime-switching mechanism endogenous. This "self-exciting" nature allows the model to directly capture nonlinearities driven by the internal dynamics of the process.

A SETAR(1) model with two regimes *(L,H)* and one lag can be written as:

$$\widetilde{p}_t = \left(\gamma_L + \delta_L \widetilde{p}_{t-1} + \nu_{L,t-1}\right) 1 \left(\widetilde{p}_{t-1} \le \theta\right) + \left(\gamma_H + \delta_H \widetilde{p}_{t-1} + \nu_{H,t-1}\right) 1 \left(\widetilde{p}_{t-1} > \theta\right) \tag{3}$$

where 1 is the indicator function and is equal to 1 if the condition is true and is zero otherwise. As can be observed, having one threshold turns the linear AR(1) process into a piecewise linear model containing two regimes, where each regime is described as a unique AR(1) process. These regimes are referred to as 'high' (H) and 'low' (L) based on the value of the forcing variable, the absolute value of the price differentials. If the price differential exceeds the thresholds, we denote the regime to be 'high' and the regime is denoted as 'low' if the price differential falls within the symmetric threshold values. In the empirical analysis, we identify the number of thresholds through a test proposed by Hansen (1999) that examines linearity against thresholds, and the SETAR(1) model against the SETAR(2) model. Specifically, we test the null hypothesis of no threshold against one threshold, no threshold against two thresholds, and one threshold against two thresholds. The value of the threshold is determined using a grid search across all possible threshold values as defined by the data.

We further extend our analysis by incorporating a bivariate Threshold Vector Error Correction (TVEC) model. While the TVEC model is not the primary focus of our study, it provides a more comprehensive framework for analyzing price dynamics by jointly modeling interactions between the two products. The first step in the model specification involves testing the statistical properties of the price series, specifically testing for stationarity and cointegration. A stationary series is characterized by statistical properties, such as mean and variance, that remain time-invariant, and nonstationary series exhibit trends or irregular fluctuations over time. We apply the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1979) to check the presence of a unit root. The lag length is selected using modified information criteria by Ng and Perron (2001). We also include the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test (Kwiatkowski et al., 1992) for stationarity, recognizing the potential issue of size distortion in the ADF test. The null hypothesis of the ADF test is that the series has a unit root. Rejection of the null hypothesis indicates that the series is stationary. The null hypothesis of the KPSS test is that the time series is stationary, therefore, rejecting the null indicates the presence of a unit root. If a nonstationary process becomes stationary after the first difference, then this process is denoted as *integrated of order one*, or *I(1)*.

If the prices are confirmed to be *I(1)*, the next step is to check for cointegration. The concept of cointegration was introduced by Engle and Granger in 1987. Given a system containing two or more time series, if all components in the system are I(d) and there exists a linear combination of them that is stationary, then we call these time series cointegrated even if the individual time series are nonstationary. Engle and Granger (1987) proposed a two-step method to test for cointegration between two non-stationary time series. The first step is to regress one variable on the other with OLS and obtain the residuals, and the second step of the procedure is to test residuals for stationarity. If the residual is stationary, then cointegration is confirmed. Taking Eq. 1 as an example, if u_t is stationary, then p_t^1 and p_t^2 are cointegrated.

Engle and Granger (1987) further introduced the Granger representation theorem, which states that if a system of time series is cointegrated, then there exists a Vector Error Correction (VEC) representation. The application of error correction mechanisms to econometrics was first introduced in the 1950s (see, Phillips, 1958). Thereon, these models have received great interest in economic empirical applications (see works of Davidson et al., 1978; Dawson, 1981; Salmon, 1982). Consider the following VEC representation with a single lag for two variables that are both integrated of order one,

$$\begin{bmatrix} \Delta p_t^1 \\ \Delta p_t^2 \end{bmatrix} = \begin{bmatrix} c_1 \\ c_2 \end{bmatrix} + \begin{bmatrix} a_1 \\ a_2 \end{bmatrix} u_{t-1} + \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix} \begin{bmatrix} \Delta p_{t-1}^1 \\ \Delta p_{t-1}^2 \end{bmatrix} + \begin{bmatrix} \varepsilon_t^1 \\ \varepsilon_t^2 \end{bmatrix} \tag{4}$$

where $u_{t-1} = p_{t-1}^1 - \beta p_{t-1}^2$ is the error correction term from Eq. 1. The coefficient vector a in front of the error correction term estimates the adjustment to deviations from the long-run equilibrium.

It should be pointed out that the Engle-Granger procedure assumes the cointegration relationship, or the adjustment of short-run dynamics toward the long-run equilibrium, is linear, symmetric, and instantaneous (Balke and Fomby, 1997). However, it is possible that the adjustment does not occur in every period. For example, in the context of market integration, a no-arbitrage condition may exist when the price difference is less than the transaction costs. Balke and Fomby (1997) describe this phenomenon in which economic agents only start the adjustment process when the deviation from the equilibrium exceeds a certain threshold as *"discrete adjustment"*. In the event of lumber product substitutability, the relationship between SPF and SYP would theoretically change from substitutes to weakly related goods only when their price differential exceeds a certain threshold value. Balke and Fomby

(1997) introduced the concept of threshold cointegration that combines the idea of cointegration with nonlinear threshold modeling. They showed that standard tests for unit roots and cointegration are valid under threshold cointegration.[7]

The VEC model in Eq. 4 can be generalized to a Threshold VEC (TVEC) model where the adjustment vector a changes depending on the state of the error correction term (ECT) (Goodwin and Holt, 1999; Goodwin and Piggott, 2001; Ramsey et al., 2021),

$$\begin{bmatrix} \Delta p_t^1 \\ \Delta p_t^2 \end{bmatrix} = \begin{bmatrix} c_1 \\ c_2 \end{bmatrix} + \begin{bmatrix} a_1^L \\ a_2^L \end{bmatrix} u_{t-1}^L + \begin{bmatrix} a_1^H \\ a_2^H \end{bmatrix} u_{t-1}^H + \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix} \begin{bmatrix} \Delta p_{t-1}^1 \\ \Delta p_{t-1}^2 \end{bmatrix} + \begin{bmatrix} \grave{o}_t^1 \\ \grave{o}_t^2 \end{bmatrix} \tag{5}$$

where the ECT is split into two regimes, 'high' and 'low.' The low regime occurs when the ECT is below the threshold, reflecting a relatively small deviation from the long-run equilibrium and a stronger market integration. Conversely, the high regime occurs when the ECT exceeds the threshold value, representing a larger deviation from equilibrium and a lower degree of market integration. This system-based approach to the paired-price system allows for the estimation of individual price responses to shocks. Hansen and Seo (2002) derived a maximum-likelihood (ML) estimator that estimates the threshold and cointegration parameters through a two-dimensional grid search. In addition, Hansen and Seo (2002) suggested a sup-LM cointegration test for the null of the linear VEC model against a TVEC model. This test was applied to ensure the presence of threshold cointegration.

Impulse response analyses offer a visual interpretation of the dynamic price patterns in response to exogenous shocks and the overall model dynamics. To examine the response of the price parity relationship and individual prices within the softwood lumber industry, we conduct both standard (orthogonalized) Impulse Response Functions (IRFs), and the Generalized Impulse Response Functions (GIRFs) following the steps outlined by Koop et al. (1996). If responses diminish quickly after an initial shock, then shocks are transitory and markets are closely related, indicating that the lumber products are substitutes. Conversely, persistent effects of shocks indicate a weaker relationship among the markets in that it takes longer for shocks to the differential to dissipate.

The standard IRF illustrates how an exogenous shock at time t impacts the process at time $t+h$ and, therefore, measures the persistent effect of shocks on the price series. An IRF for a process y can be defined as:

$$IRF_y(h,\rho,\omega_{t-1}) = E[y_{t+h}|\grave{o}_t=\rho,\grave{o}_{t+h}=0,\omega_{t-1}] - E[y_{t+h}|\grave{o}_t=0,\grave{o}_{t+h}=0,\omega_{t-1}] \tag{6}$$

where ω_{t-1} is the observed history of the process, ρ is the size of the shock at time t, and $\grave{o}_t$ is the error term. One important thing to note is that for the standard IRF, $\grave{o}_{t+h} = 0$, that is, no further shocks are assumed to affect the system. Using parameters from each regime of the threshold models, we perform regime-specific impulse responses to observe distinctive lumber price behaviors under different regimes.

Koop et al. (1996) pointed out that the standard IRF is less useful when applied to nonlinear models because the response may differ according to the size, sign, and timing of a shock. In the case of the threshold model, the impacts of the shock for the standard IRF would differ depending on the size and starting regime of ω_{t-1}. Therefore, they introduced GIRFs that allow for the examination of responses not only to one standard deviation shock in a single variable but also to shocks in multiple previously observed variables simultaneously. Consider the following GIRF on process y:

[7] Balke and Fomby (1997) also noted that these tests may lack power in the presence of threshold effects.

$$GIRF_y(h,\rho,\omega_{t-1}) = E[y_{t+h}|\grave{o}_t=\rho,\omega_{t-1}] - E[y_{t+h}|\omega_{t-1}] \tag{7}$$

With GIRF, instead of using y_{t+h} without shock ρ as the benchmark, the compared realization of y_{t+h} is an average of future shocks. For the calculation of the realized GIRFs, we follow Koop et al.'s suggestion and draw random bootstrap samples of the estimated error term for the shock ρ and observed price for ω_{t-1}. Similar to the standard IRFs, we consider the regime-specific GIRFs for both responses of the price parity relationship and individual lumber prices.

4. Data and model specification

Weekly lumber prices from January 1995 to April 2020 were collected from a private commodity price reporting agency *Fastmarkets*. We intentionally truncated the dataset to exclude the period following extreme price volatility in the lumber industry during the COVID-19 pandemic. The surge in softwood lumber prices beginning in May 2020 can be attributed to supply chain disruptions, structural shifts in the residential construction and labor markets, and temporary policy interventions—factors that were not present during other periods. Therefore, by excluding these prices, we can analyze the relationship between domestically produced and imported lumber without the influence of abnormal exogenous shocks in the economy resulting from the pandemic. We focus on Western Spruce-Pine-Fir (SPF) shipped from Canada,[8] and Southern Yellow Pine (SYP) produced in the west side of the Southern U.S.[9] The selected lumber dimensions for both products are 2×10 (inches) random lengths with a lumber grade of 2. All price series are illustrated in Fig. 1, with all prices expressed in U.S. dollars per thousand board feet (mbf). SYP prices are free-on-board (FOB) at the mills and SPF prices include duties or export taxes when applicable.

SPF is the largest imported lumber species group from Canada, constituting 86 % of total Canadian softwood lumber imports; While SPF is also produced in the U.S., it only accounts for 5.4 % of total softwood production (Shook et al., 2007). SYP, on the other hand, is the largest single species of lumber produced in the U.S., and no SYP is commercially produced in Canada (Shook et al., 2007). It is worth noting that lumber products used in the analysis are kiln-dried and untreated, indicating that they have not undergone chemical processing to resist rot, moisture, and decay. Treated lumber is commonly used for outdoor purposes whereas untreated lumber is primarily used for structural uses in residential constructions. As noted by Nagubadi et al. (2004), treated and untreated lumber products may be complements, since they are often used in the same construction site, but for entirely different uses.

The price series collected comprises 1321 weeks.[10] We observe that the means of SPF and SYP are close, with SYP prices showing higher volatility. As shown in Fig. 1, we observe that in general, SPF and SYP follow similar patterns, which is an indication of possible cointegration and may be indicative of substitutability.

Prices are transformed to logarithmic forms for the rest of the analysis, which allows us to interpret price differentials and marginal effects in percentage terms. We conduct stationarity tests for individual lumber prices and price differentials using the ADF and KPSS tests. Test statistics are presented in Table 1, and p-values are included in brackets. Rejecting the null hypothesis of the ADF test implies stationarity while rejecting the null hypothesis of the KPSS test implies the presence of a unit root.

[8] Western SPF: shipped from Prince George, British Columbia, Canada. Duties/export taxes included when applicable.

[9] Westside SYP: mills in Texas, Arkansas, Oklahoma, and Louisiana west of the Mississippi River. Prices are FOB at the mills.

[10] From January 01, 1995, to April 24, 2020, the average SYP price in the U.S. was \$351.82/mbf, with a standard deviation of \$82.59/mbf, and prices ranging from \$193.00 to \$596.00/mbf. In comparison, the average price of SPF from Canada was \$353.29/mbf, with a standard deviation of \$70.09/mbf, and prices ranging from \$145.00 to \$601.00/mbf.

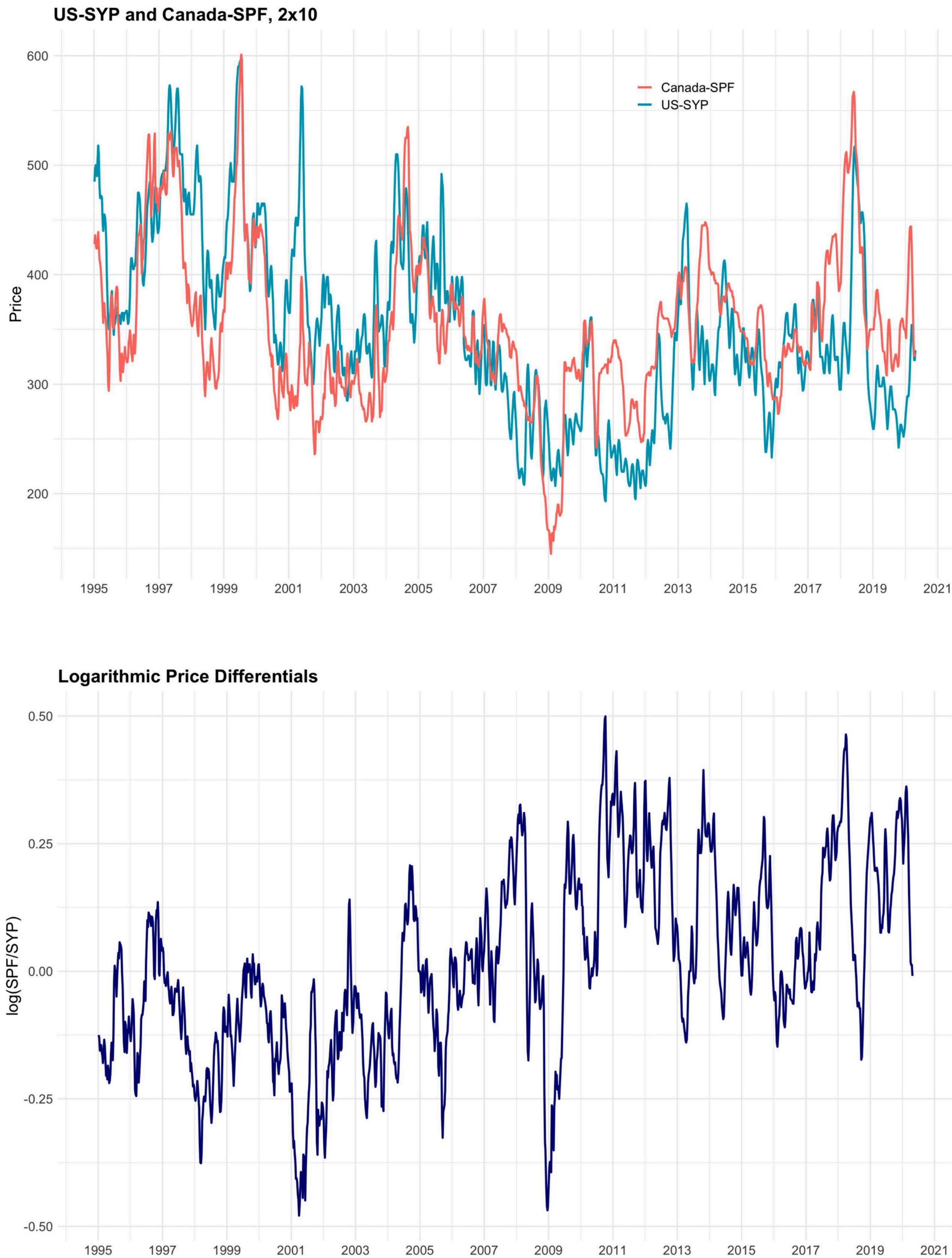


Fig. 1. Weekly Price Movements and Price Differentials of Canada SPF and U.S. SYP January 6, 1995, to April 24, 2020 (No. Observation: 1321).

Table 1
Unit root test for individual price series and price differentials.

		Level Data		First-Differenced Data	
		ADF Test[1]	KPSS Level[2, 3]	ADF Test	KPSS Level
Individual prices	US-SYP	−0.256 [0.594]	6.317 [0.010]	−19.54 [0.000]	0.022 [0.100]
	CA-SPF	−0.106 [0.647]	1.137 [0.010]	−14.41 [0.000]	0.021 [0.100]
Abs(price diff)	Canada-US	−2.953 [0.000]	0.334 [0.100]	–	–

Brackets contain *p*-values.
[1] Null hypothesis of the ADF test is the series contains a unit root. vs. Alternative: The series is stationary.
[2] Null hypothesis of the KPSS test is: The series is stationary. vs. Alternative: The series contains a unit root.
[3] The critical value at 10 % is 0.347, and the critical value at 1 % is 0.739. Computed statistics are outside the table of critical values listed in Kwiatkowski (1992), therefore, actual p-values are less than 1 % for the level data and greater than 10 % for differenced data.

Table 2
Test for Cointegration, Linearity, and Linearity Against Threshold.

Model/Test	Statistics Description	Test Statistics
Price transmission	α	2.529 [0.000]
	β	0.568 [0.000]
	Adjusted R^2	0.437
EG two-step cointegration test	ADF test of error term	−4.425 [0.000]
Hansen and Seo's test[1]	Linear vs. Threshold cointegration	35.440 [0.024]
Linearity tests	White[2]	12.202 [0.002]
	Keenan[3]	1.629 [0.202]
	Tsay[4]	3.605 [0.000]
	Chan[5]	94.331 [0.000]
Hansen's test: Linearity vs. Threshold[6]	AR vs. SETAR(1)	110.732 [0.000]
	AR vs. SETAR(2)	127.388 [0.000]
	SETAR(1) vs. SETAR(2)	17.562 [0.101]

Brackets contain p-values.
[1] Hansen and Seo (2002) test for two-regime threshold cointegration in VECMs. N_0: Linear VECM. *P*-values are calculated from 1000 bootstrap replications.
[2] White (1993) Neural Network Test. N_0: The time series is linear in mean.
[3] Keenan (1985) One-Degree Test. N_0: The time series follows some AR process.
[4] Tsay (1986) Nonlinearity Test. N_0: The true model is an AR process.
[5] Chan (1993) Likelihood Ratio Test. N_0: The time series follows a normal AR process.
[6] Hansen (1999) test for threshold effects in AR models. If the null hypothesis of a linear AR model is rejected, then proceed to test the number of thresholds. P-values are calculated from 1000 bootstrap replications.

The results show that both lumber prices are nonstationary as we fail to reject the null hypothesis of the ADF test. This is consistent with the results of the KPSS test, where the null hypothesis is strongly rejected for both lumber prices, indicating the price series contains a unit root. In addition, both ADF and KPSS indicate that prices and price differentials become stationary after the first difference, suggesting that they are integrated of order one. Next, we assess the stationary properties of the absolute value of the price differentials, where the test statistics are presented in Table 1. The results of both the ADF and KPSS tests indicate that the price differentials are stationary. We then check for cointegration.

The process of testing for cointegration begins with Engle-Granger's two-step approach, by first estimating a linear regression between each paired price, followed by testing for stationarity of the estimated residuals. As suggested by Engle and Granger, if the residuals are stationary, then the two prices are considered cointegrated. Cointegration test results are presented in Table 2. The null hypothesis of the ADF test for estimated residuals for each paired price is rejected at least at the 1 % level. This suggests strong evidence supporting cointegration between Canadian and U.S. softwood lumber prices. We further apply Hansen and Seo (2002) test of linear cointegration against threshold cointegration. *P*-values are calculated from 1000 bootstrap replications. The null hypothesis is rejected, indicating the presence of threshold cointegration.

We proceed with the model specification procedure, conducting tests to identify nonlinear behaviors in the price differentials. Linearity test results are described in Table 2. Evidence of nonlinear patterns is mixed but generally favors a nonlinear specification, with three out of the four tests rejecting the null hypothesis of linearity at the 1 % level. With the confirmation of nonlinearity, we now determine the number of thresholds by conducting three tests as proposed by Hansen (1999) test for linear Autoregressive (AR) model against a Self-Exciting Threshold Autoregressive (SETAR) model. The three tests are as follows, with test results listed in Table 2: no threshold against one threshold (AR vs. SETAR(1)), no threshold against two thresholds (AR vs. SETAR(2)), and one threshold against two thresholds (SETAR(1) vs. SETAR(2)). The results favor one threshold. To determine the number of lags included in each model, we compute Akaike information criterion (AIC) and Bayesian information criterion (BIC) for up to five lags. The computed criteria are listed in Table A.1. We find the SETAR(1) model with four lags minimized both the AIC and BIC criteria, while the TVEC model with three lags similarly minimized both AIC and BIC. Therefore, the optimal models selected are SETAR(1) with four lags and a TVEC(3).

5. Empirical results

The first step of the analysis is to estimate the standard linear price transmission model described in Eq. 1, where we regress the logarithmic price of SPF on SYP. Table 2 contains the estimated parameters of the model, where β is the price transmission elasticity. Note that $\beta = 1$ implies perfect integration. In the conventional spatial market integration analysis, price transmission occurs when speculators transfer goods from one market to another when arbitrage opportunities arise. However, in this study, price transmission behavior is also driven by substitution in the consumption of different lumber species. The slope coefficient β is 0.568, suggesting markets for the two lumber products are not very well integrated. If the alternative products were perfect substitutes, we would expect a price elasticity of one. Although the magnitude of substitution and relationships between lumber prices cannot be precisely depicted by this simple price transmission model, the positive and significant correspondence between prices affirms the price linkage among Canadian and U.S. lumber prices, albeit with a less-than-perfect degree of integration.

We continue the analysis with univariate model estimations of the price differentials. Since nonlinearity is confirmed, we apply a Self-Exciting Threshold Autoregressive (SETAR) model with a single threshold. The threshold θ is searched over the $[\theta_{15\%}, \theta_{85\%}]$ interval, where $\theta_{10\%}$ corresponds to the 10 % quantile of the threshold value.[11] The estimated threshold value of the SETAR model is 0.034. Since the forcing variable in this analysis is the absolute value of the price differentials, a threshold value of 0.034 establishes a natural band between − 0.034 and + 0.034. This creates two regimes: the 'low' regime, corresponding to observations within the band, and the 'high' regime, corresponding to those outside the band. The low regime comprises

[11] We explored lower quantiles, such as the 5th and 10th percentiles, but found that they yielded the same estimate as the 15th percentile.

Table 3
Comparison of Price Transmission Elasticity.

	Intercept α		Slope β		
Regimes	Estimate	Std. err	Estimate	Std. err	Adjusted R^2
Within Band	0.100	0.043	0.983	0.007	0.989
Outside Band	0.052	0.117	0.519	0.020	0.376

16.1 % (212 observations) of the dataset. Based on our hypothesis, under the low regime, when the price parity between SPF and SYP is less than 3.34 %, the two lumbers act as substitutes. On the other hand, the high regime contains 83.9 % (1105 observations) of the dataset. It occurs when SPF prices deviate from SYP prices by more than 3.34 %, either above or below. In this regime, we hypothesize that the price linkage becomes relatively weak, with a lower degree of market integration, thereby implying less substitutability between the two lumber species.

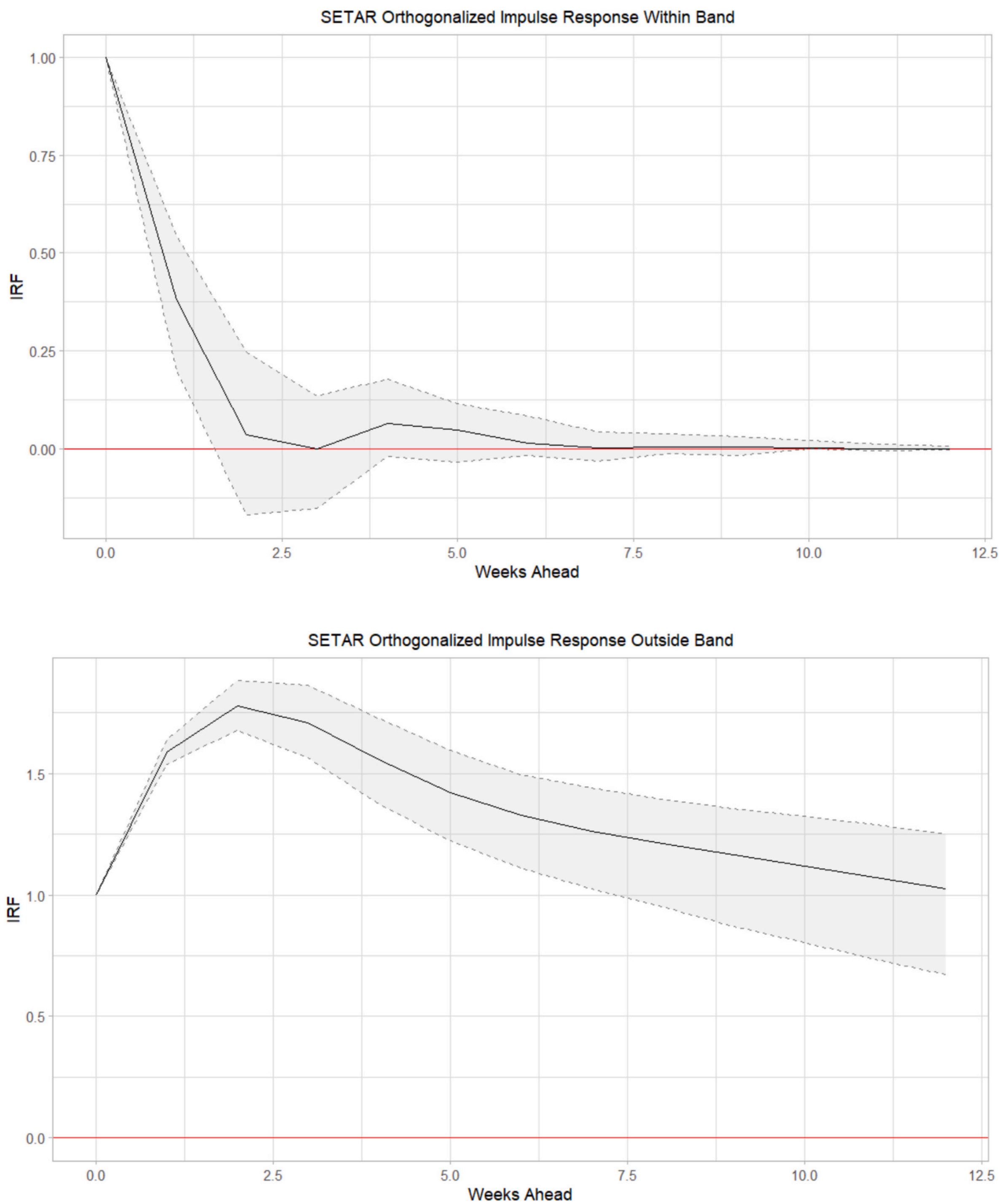
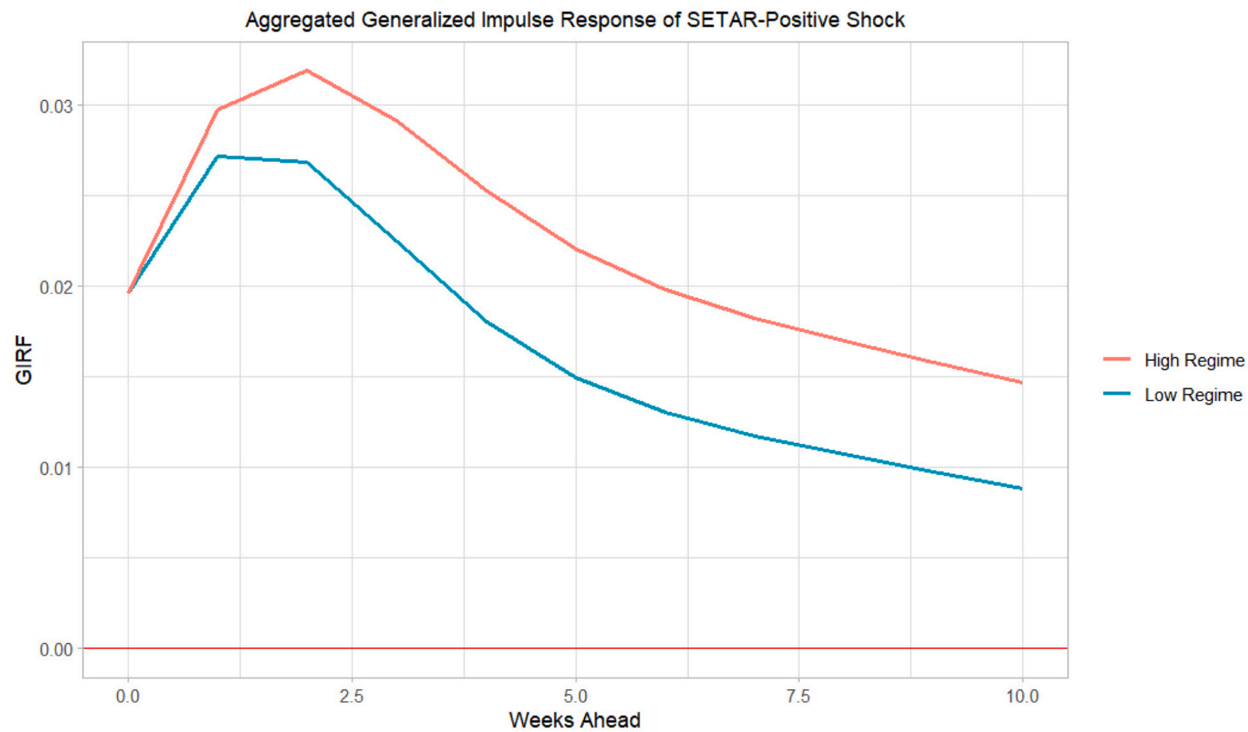


Fig. 2. Regime-Specific Orthogonalized Impulse Response of SETAR Model.
This figure presents the regime-specific orthogonalized impulse response of the SETAR model. The top panel presents the impulse response in the low regime, and the bottom panel presents the impulse response in the high regime. IRFs are generated based on a SETAR(1) model. The data frequency is weekly. All variables are transformed into natural logs. The shaded areas represent 95 % bootstrapped error band.

(a) Positive Shock



(b) Negative Shock

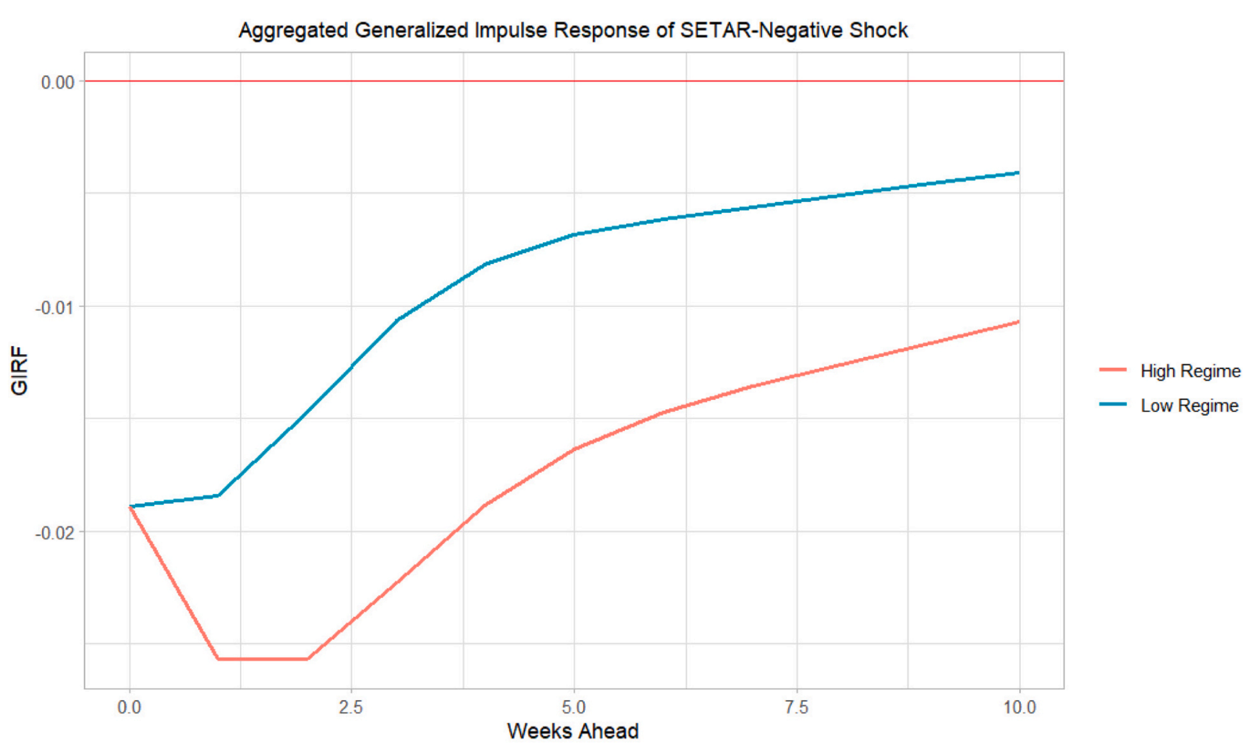


Fig. 3. Aggregated Generalized Impulse Response of SETAR Model.

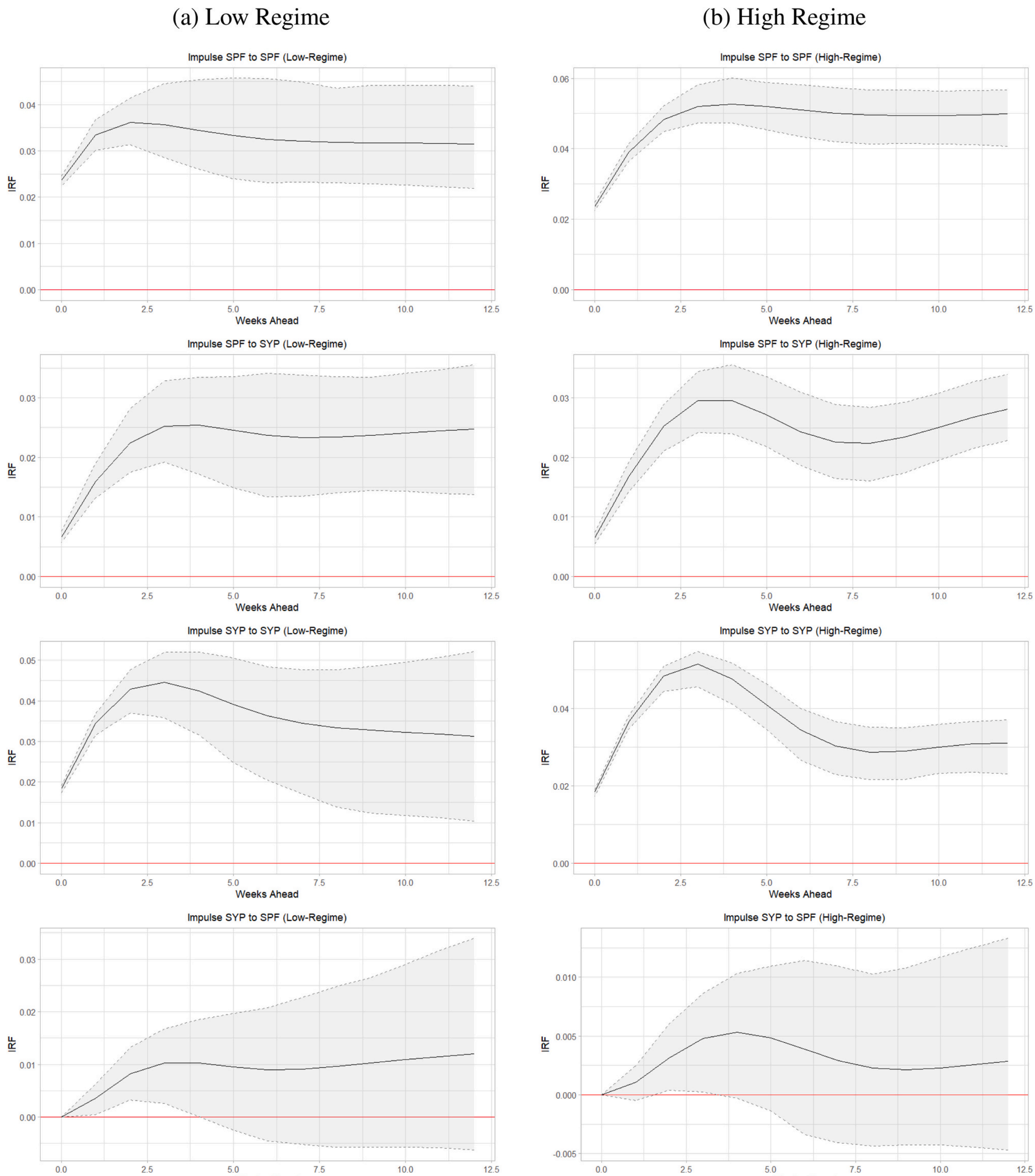
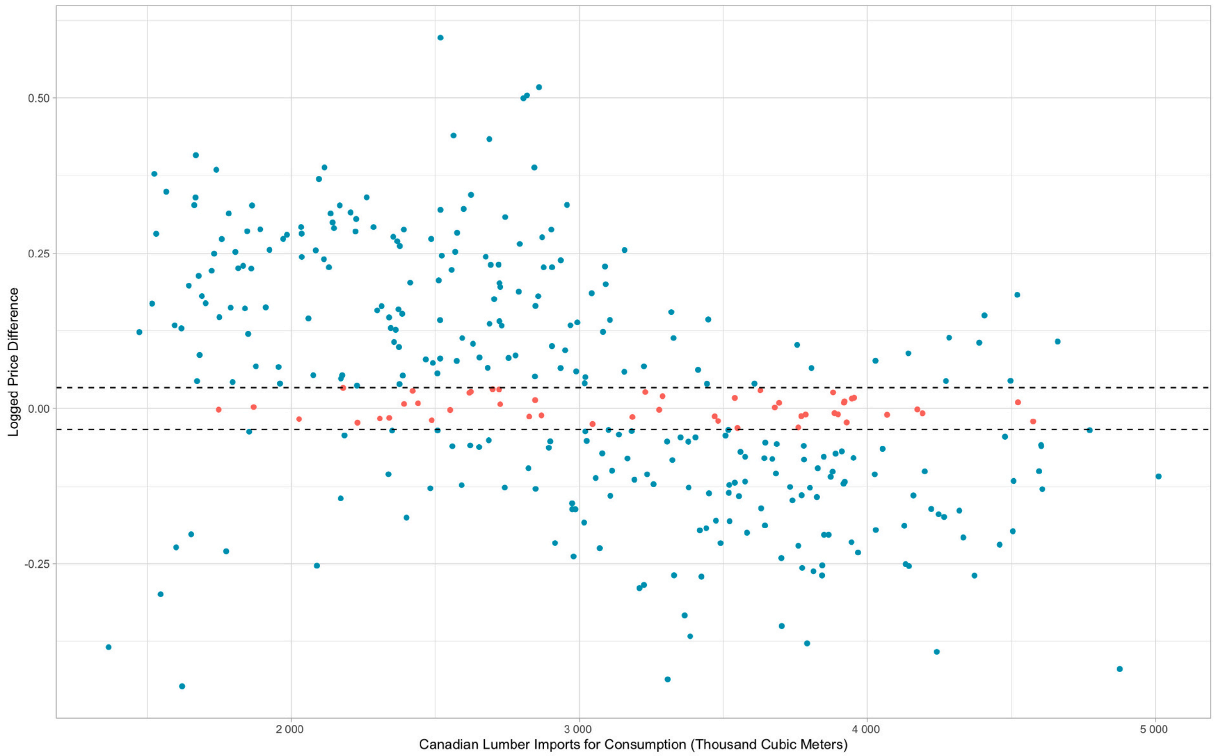


Fig. 4. Regime Specific Orthogonalized Impulse Response of TVEC Model.
Note: This figure presents the regime-specific orthogonalized impulse response of the TVEC model. Left panels under (a) present impulse responses in the low regime, and right panels under (b) present impulse responses in the high regime. IRFs are generated based on a TVEC(3) model. The data frequency is weekly. All variables are transformed into natural logs. The shaded areas represent 95 % bootstrapped error band.

In such cases, we would expect consumers to select the lumber product with the lower price.

Once the natural band is defined, we estimate the standard linear price transmission model for each regime separately. The estimates of the linear price transmission model are presented in Table 3. Notably, the price transmission elasticity within the band is 0.983, implying almost perfect integration, while the estimated elasticity outside the band is only 0.519. This significant difference between the two regimes indicates a much higher degree of market integration within the band compared to outside it. Moreover, the higher adjusted R^2 in the low regime highlights the strong linkage between the two products, in contrast to the weaker linkage reflected by the lower adjusted R^2 in the high regime. This simple linear analysis offers a straightforward comparison of the degree of market integration across regimes, supporting



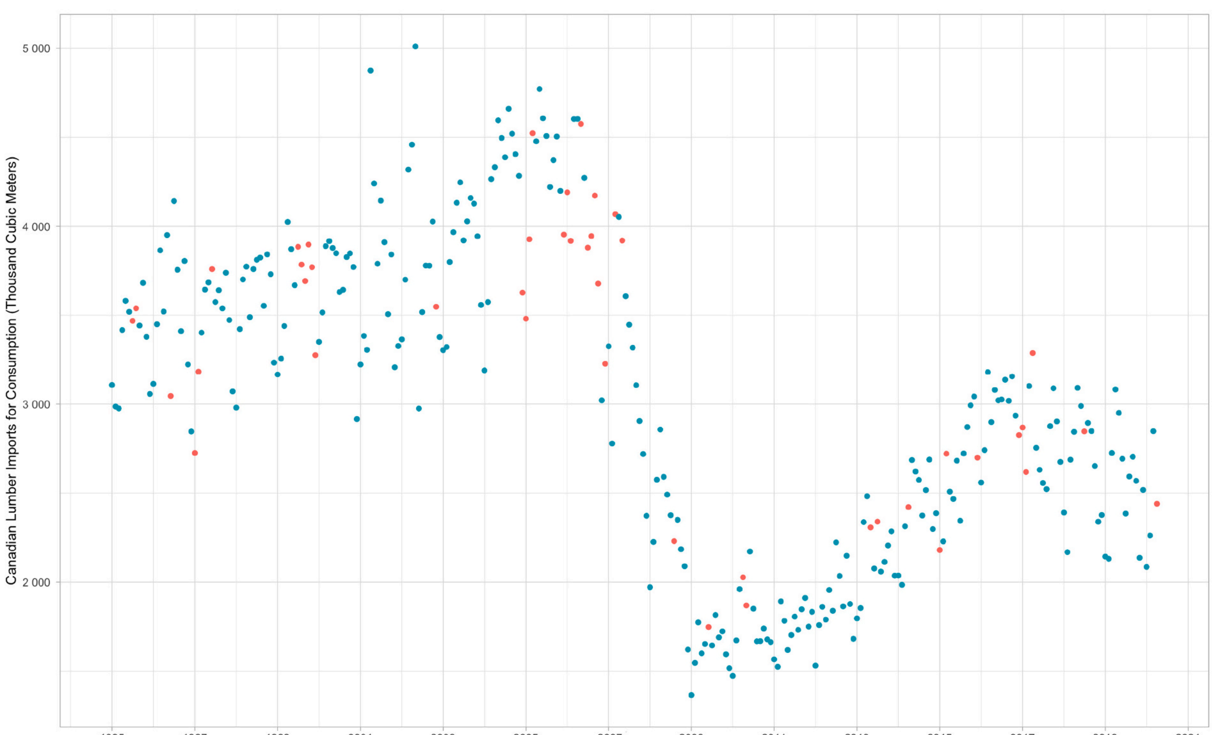


Fig. 5. Canadian Lumber Imports for Consumption in the U.S. (1995–2020).
Note: The top figure displays a scatter plot of logged price differentials against Canadian lumber import volumes, with the natural band of $+/-$ 0.034 estimated from the SETAR model. The bottom figure plots the monthly Canadian lumber import volumes from January 1995 to April 2020, where the pink dots indicate periods during which the price differentials fall within the estimated natural band. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

our hypothesis regarding substitutability.

Our hypothesis of lumber substitutability is further supported by the results from impulse response analyses. The calculation of IRFs is less trivial for threshold models compared to linear models, as responses depend on the state of the regime. We first consider regime-specific IRFs, constructed using parameters estimated from the SETAR model. The differences in persistence across regimes are evident in the regime-specific IRFs, as depicted in Fig. 2. The top figure illustrates the IRF in the low regime, where the effects of shocks dissipate more rapidly than in the high regime as shown in the bottom figure, suggesting that shocks are transitory. A fast convergence to equilibrium following a shock indicates a high degree of market integration. In the context of cross-commodity price transmission, such convergence in price differentials suggests that consumers are responsive to deviations in price parity, engaging in substitution behaviors that bring the two prices closer. This rapid convergence in the IRF reflects the effectiveness of market integration in correcting price discrepancies. Conversely, the IRF in the high regime indicates significantly higher persistence, indicating a slower adjustment to shocks, which suggests a lower degree of substitutability. These IRF results are consistent with our hypothesis that closer price proximity leads to greater substitution between the two lumber products.

One drawback of the standard IRF is that it only reflects the effect of a one-time positive shock at time $t+h$, compared to innovation at time $t = 0$. As pointed out by Koop et al. (1996), when nonlinearity is present, the nature of the responses may also depend on the direction and timing of the shocks. Therefore, we include the Generalized Impulse Response Functions (GIRFs) for the SETAR model, following the unified approach suggested by Koop et al. (1996). They recommend constructing GIRFs using different sets of history and shocks, as these two values jointly determine the response paths. We implement Monte Carlo simulations with 500 bootstrapped histories drawn from the observed data and 500 shocks drawn from the estimated residuals. GIRFs are then computed using 100 replications from the estimated based line and shocked path as described in Eq. 7.[12]

It is most meaningful to consider the simulated GIRF replications based on the regime and direction of the shocks. We divide the replicated GIRFs into four groups to explore density plots and aggregated responses. First, we obtain the regime-specific GIRFs by separating the replications based on whether simulations were started in the high or low regimes. We further separate the regime-specific GIRFs given the direction of shocks, which allows us to observe if price differentials respond symmetrically to positive and negative shocks. We then calculate the aggregated path of GIRFs.[13] As presented in Fig. 3, GIRFs are aggregated by regimes and then by the direction of the shocks. From the aggregated GIRF plots, we make the following observations. First, shocks in the high regime show a significantly larger magnitude and higher persistence than those in the low regime, which holds true for both positive and negative shocks. This finding aligns with findings from regime-specific IRFs. Second, we find relatively symmetric responses in the high regime to positive and negative exogenous shocks, while in the low regime, positive shocks induce a slightly larger adjustment than negative shocks. In summary, the GIRF results from the SETAR model are consistent with our hypothesis, indicating that lumber prices in the low regime suggest a higher degree of substitution compared to those in the high regime.

While it is useful to investigate the univariate model based on the price parity relationship, a bivariate approach can provide a more comprehensive and nuanced analysis of specific changes for each price within the system. Since threshold cointegration is identified in the pairwise price system, a Threshold Vector Error Correction (TVEC) model is applied. The TVEC model is estimated in a similar fashion compared to the univariate SETAR model with one threshold and, therefore, the model identifies two regimes. In TVEC models, the regime-switching depends jointly on the cointegrating vector $(1, \beta)$ and the threshold parameter θ. A grid search is conducted over each value of (β, θ) within the 15 % and 85 % quantiles, and the pair minimizing the sum of residual squared errors is selected. The estimated cointegration vector is (1, −1.004) with a selected threshold value of −0.175. The forcing variable of the TVEC model is the error correction term (ECT). Therefore, the low regime is triggered when $ECT = SPF_t - \beta \times SYP_t$ is below the threshold, i.e., $SPF_t \leq 1.004 \times SYP_t - 0.175$. Since the estimated cointegration vector is close to $(1, -1)$, the ECT can be interpreted as price differentials. The low regime contains 18.8 % of the dataset (248 observations), indicating periods where the deviation from equilibrium is small, while the high regime, containing 81.2 % of the dataset (1069 observations), occurs when SPF prices are less than 17.5 % lower than SYP prices, reflecting larger deviations from the long-run equilibrium.

The impulse response analysis of the multivariate system provides richer information as it traces the impacts of changes in one variable on others in the system. For the TVEC model, the regime-specific IRFs are plotted in Fig. 4. The panels on the left show impulse responses in the low regime, and the panels on the right show the responses in the high regime (values of the error correction term that are greater than the threshold). The difference in the speed of adjustment between the two regimes is apparent. The impulse responses in the high regime show no sign of convergence after 12 weeks, whereas in the low regime, the responses become insignificant after week six in most cases. The magnitude of the responses is also much larger in the high regime in both variables. Lastly, a notable difference between the high and low regimes is the width of the confidence bands. Impulse responses in the low regime exhibit wider confidence bands compared to those in the high regime. This can be attributed to the smaller number of observations in the low regime, which leads to greater variability. These results suggest similar implications as those from the univariate models. The rapid adjustments in the low regime indicate the higher degree of substitutability of the two lumber markets, while the sustained persistence in the high regime suggests weaker integration between the two lumber products.

Following a similar procedure as in the univariate case, we compute GIRFs over 100 bootstrap replications from 500 draws of the history and estimated residuals. The shock we consider is a system-wide shock, meaning residuals from both SPF and SYP are contemptuously selected. The aggregated GIRF paths are presented in Fig. A.1, confirming the varying shock persistence levels observed in the regime-specific IRFs. For both SPF and SYP, we find that an average shock in the high regime produces stronger and more persistent impacts, regardless of the direction of the shocks. In addition, we observe a faster convergence in the low regime, especially for SPF. These results are consistent with the standard IRF and findings from the SETAR model, further supporting our hypothesis of greater market integration and a higher degree of substitutability within the low regime.

Regulations on softwood lumber imports can introduce potential structural breaks in the price transmission. A particularly relevant break point to consider is the implementation of the 2006 Softwood Lumber Agreement (SLA2006) on October 12, 2006 (Zhang, 2007). We conducted the standard Chow test for a VEC model for the lumber price pair, which the null hypothesis of structural stability was rejected.[14] Recognizing that a structural break could introduce potential biases into our analysis, we segmented the dataset and applied the SETAR model

[12] This procedure requires high computational intensity. With 500 histories, 500 shocks, and 100 replications for each, amounting to 25 million simulations.

[13] GIRFs are aggregated by mean, a typical approach in empirical applications. We acknowledge that aggregation may lead to a loss of information conveyed in the response. Nevertheless, we believe the aggregated paths offer valuable insights when comparing across regimes and directions of shocks.

[14] The Chow test statistics is 12.90, and the corresponding p-value is 0.00, suggesting there is a structural break around this breakpoint.

separately to the pre-break and post-break periods. The approach allowed for an independent examination of regime-switching behavior within each period. The detailed results of this segmented analysis are provided in Appendix Fig. A.2. Note that the fitted models for the pre-break and post-break periods yield similar threshold values (0.033 and 0.027) and impulse responses across regimes. We have also fitted a VEC model for periods before and after the breakpoint, with the impulse responses presented in Appendix Fig. A.3. Again, the results are remarkably similar. These findings validate the robustness of our primary SETAR model, even in the presence of a structural break.[15]

It is potentially instructive to consider how the observations are related to the thresholds of $+/-$ 3.34 %. Fig. 5 plots the individual price differentials against imports of Canadian lumber into the U.S., taken from the U.S. Department of Commerce. Both panels identify price differentials that fall within the band denoted by the thresholds. The upper panel shows that imports tend to fall as the price differential decreases, making Canadian lumber relatively more expensive. The lower band plots imports from Canada over time, where we observe a sharp decline in Canadian lumber imports following the implementation of SLA2006, likely driven by the export charges imposed on Canadian exporters when softwood lumber prices fell below a specified reference level, as well as the volume quotas that restricted exports during periods of low prices (Zhang, 2007; Nagubadi and Zhang, 2013). The price differentials within the band (pink dots) occurred slightly more frequently when trade flows were greater, suggesting a greater degree of substitutability during this period. The diagram indicates that the change in policy had the most significant impact on imports of lumber from Canada.

Overall, our findings in the univariate price differential approach align with those from the multivariate error correction approach. With the confirmed threshold effect in both cases and the lower persistence profile in the low regime in response to a shock, the evidence supports our hypothesis that, in the low regime, SPF and SYP exhibit a much lower degree of substitution.

6. Conclusion

Trade in softwood lumber between the U.S. and Canada has significant impacts on both economies, as the lumber market is closely linked to construction, manufacturing, and labor industries. In recent years, global forest products production and consumption have experienced an upswing, driven by increased residential construction, highlighting the importance of addressing the trade dispute between the U.S. and Canada. The history of the U.S.-Canada Softwood Lumber Dispute involves a recurring pattern of trade litigation and agreements, and the determination of injury to the U.S. industry from Canadian lumber imports is a crucial part of the development of a new trade agreement. Motivated by this aspect of the trade dispute, this paper examines the substitutability between U.S. and Canadian softwood lumber species through a threshold analysis. Theoretically, if the degree of substitutability is not substantial, the implementation of trade restrictions on Canadian softwood imports is likely to have minimal impact on the U.S. market.

For lumber prices collected between 1995 and 2020, we find strong evidence supporting the presence of threshold effects between Spruce-Pine-Fir (SPF) imported from Canada and Southern Yellow Pine (SYP) produced in the U.S., which can be well characterized by models with one symmetric threshold. We examine the nonlinearity in price differentials using a univariate approach that utilizes a Self-Exciting Threshold Autoregressive (SETAR) model. The SETAR model selects a threshold band of $+/-$ 0.034. We hypothesize that when the price parity between the two lumber products falls within the threshold band, referred to as the low regime, the two products exhibit substitutability, as consumers are more inclined to substitute one for the other. Conversely, when the price differential exceeds the threshold band, the products exhibit weaker market integration and tend to behave more independently. Our hypothesis is supported by impulse response analysis results that the shock persistence in the low regime is much lower than in the high regime. Dynamics of market responses contribute important understanding to the broader discourse on trade policy in the softwood lumber industry.

While this study only focuses on two lumber species, the analysis can be easily generalized to other lumber species produced in other regions of the U.S. and Canada. Overall, we bring an important view on the substitutability of two lumber species, which is an addition to the literature on the softwood lumber dispute and can provide additional information looking into the future of the dispute. Understanding how price parities between domestic products and imports could potentially provide valuable information on future trade restrictions. For example, U.S. trade negotiators can consider price parities as a factor to determine the appropriate threshold when placing trade duties or other trade-restricting policies on Canadian lumber imports.

CRediT authorship contribution statement

Yifei Zhang: Writing – original draft, Visualization, Validation, Software, Resources, Project administration, Methodology, Investigation, Formal analysis, Data curation, Conceptualization. **Barry K. Goodwin:** Writing – review & editing, Validation, Supervision, Conceptualization, Resources.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

Data will be made available on request.

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[15] Figure 5 illustrates a significant drop in imported lumber volume during the Roman et al., 2006–2009 period. This decline likely reflects the impact of the SLA2006 after the policies were realized by the market. To ensure the robustness of our findings, in addition to using the effective date of the SLA2006 as a breakpoint, we conducted an additional analysis using December 2009 as an alternative breakpoint and fitted a VEC model to the data before and after this date. The results, presented in Appendix Figure A.4, show that the pre- and post-2009 analyses yield very similar outcomes, suggesting that potential structural breaks do not significantly affect our primary conclusions.

Appendix A

Table A.1
Model Specification: Information Criteria.

	AIC					BIC				
	Lag = 1	Lag = 2	Lag = 3	Lag = 4	Lag = 5	Lag = 1	Lag = 2	Lag = 3	Lag = 4	Lag = 5
CA-US (N = 1321)										
SETAR	−9141.84	−9569.38	−9600.94	−9603.66	−9600.21	−9000.18	−9495.03	−9527.01	−9531.22	−9522.71
TVEC	−20,029.64	−20,369.24	−20,351.69	−20,334.27	−20,314.31	−19,894.83	−20,172.25	−20,092.54	−20,012.97	−19,930.87

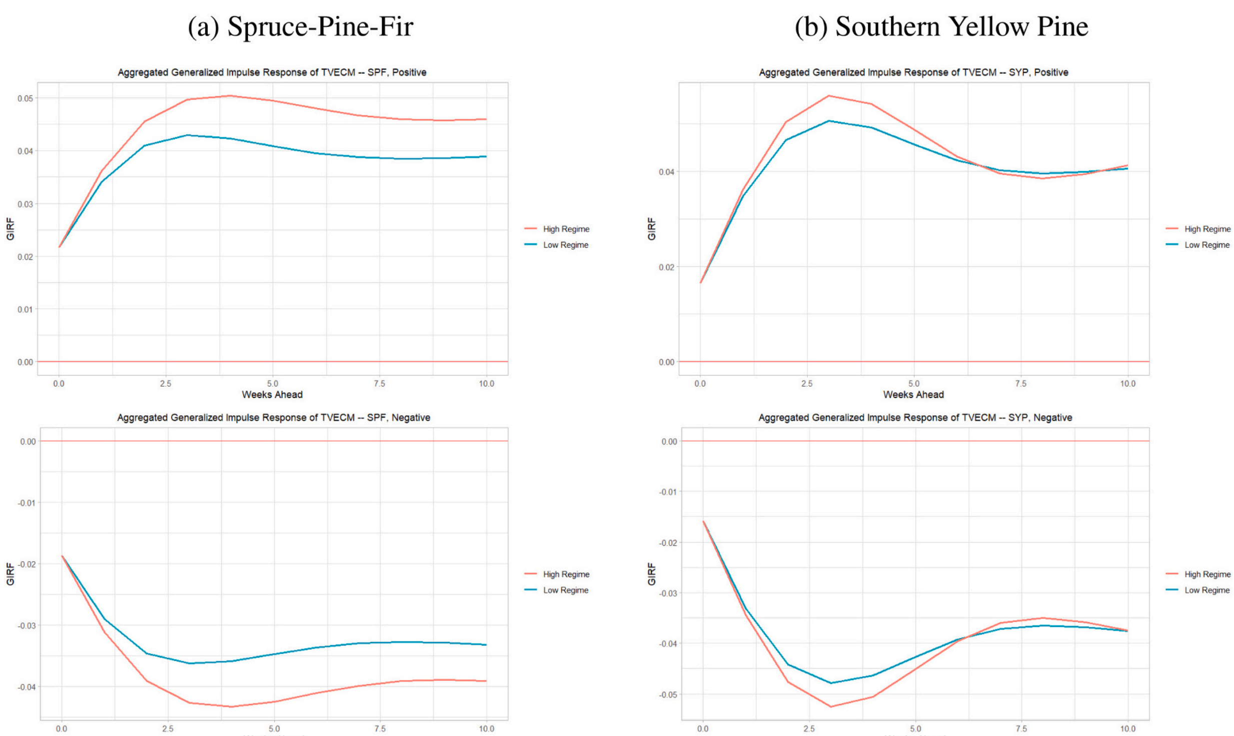

Fig. A.1. Aggregated Generalized Impulse Response of TVEC Model.
Note: This figure presents the aggregated generalized impulse responses (GIRFs) TVEC. Left panels under (a) present GIRFs for SPF in response to a positive and negative economy-wide shock, and right panels under (b) present GIRFs for SYP in response to a positive and negative economy-wide shock. IRFs are generated based on a TVEC(3) model. The data frequency is weekly. All variables are transformed into natural logs.

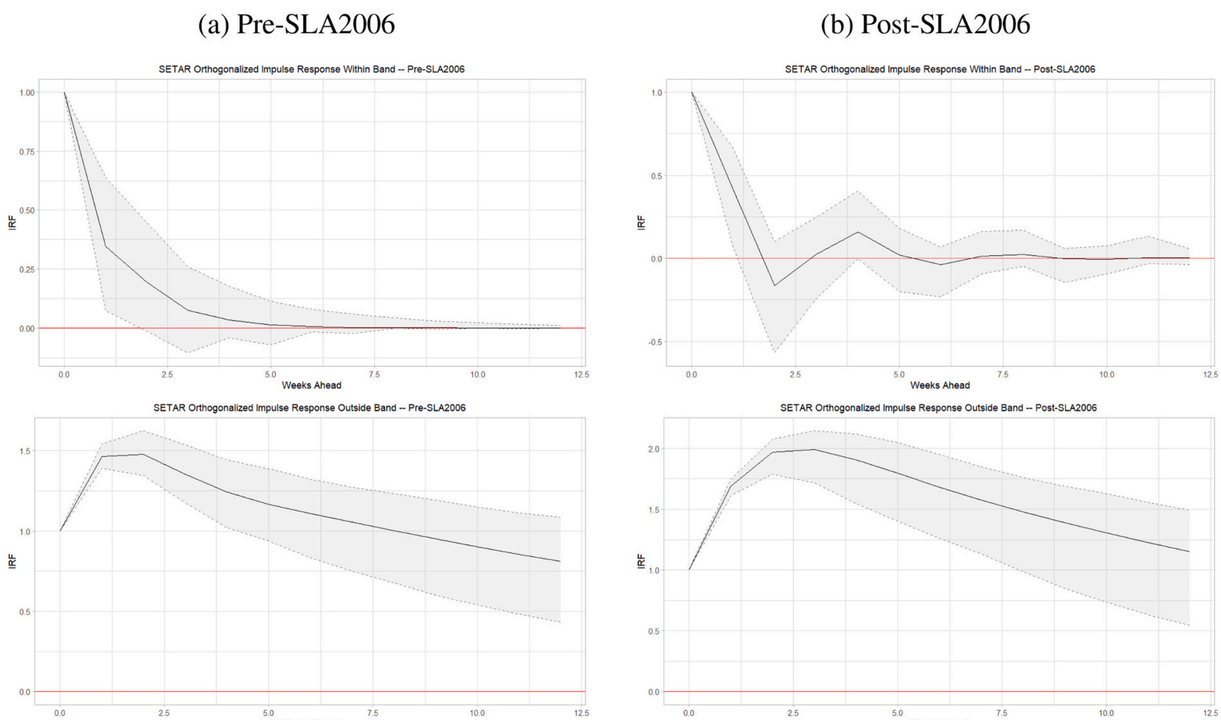

Fig. A.2. Regime Specific Orthogonalized Impulse Response of TVEC Model.
Note: This figure presents the regime-specific orthogonalized impulse responses of SETAR. Left panels under (a) present impulse responses before the effect date of the SLA2006, and right panels under (b) present impulse responses after the effect date of the SLA2006. IRFs are generated based on a SETAR(1) model. The data frequency is weekly. All variables are transformed into natural logs. The shaded areas represent 95 % bootstrapped error band.

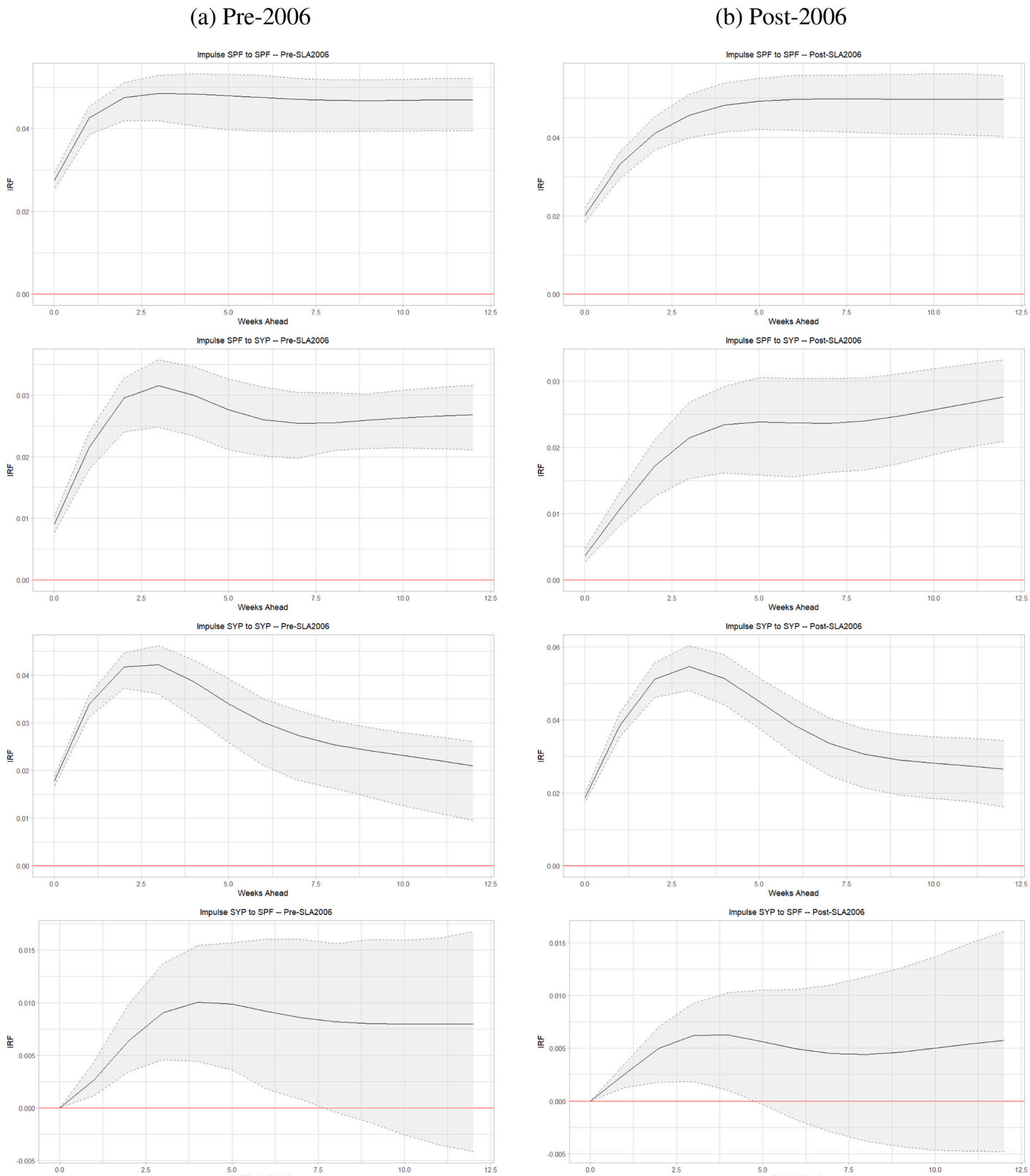

Fig. A.3. Orthogonalized Impulse Response of VEC Model – Pre-and Post-SLA2006.
Note: This figure presents the regime-specific orthogonalized impulse responses of VEC models using segmented data on 10/12/2006, the effective date of SLA2006. Left panels under (a) present impulse responses before the break date, and right panels under (b) present impulse responses after the break date. IRFs are generated based on a VEC(2) model. The data frequency is weekly. All variables are transformed into natural logs. The shaded areas represent 95 % bootstrapped error band.

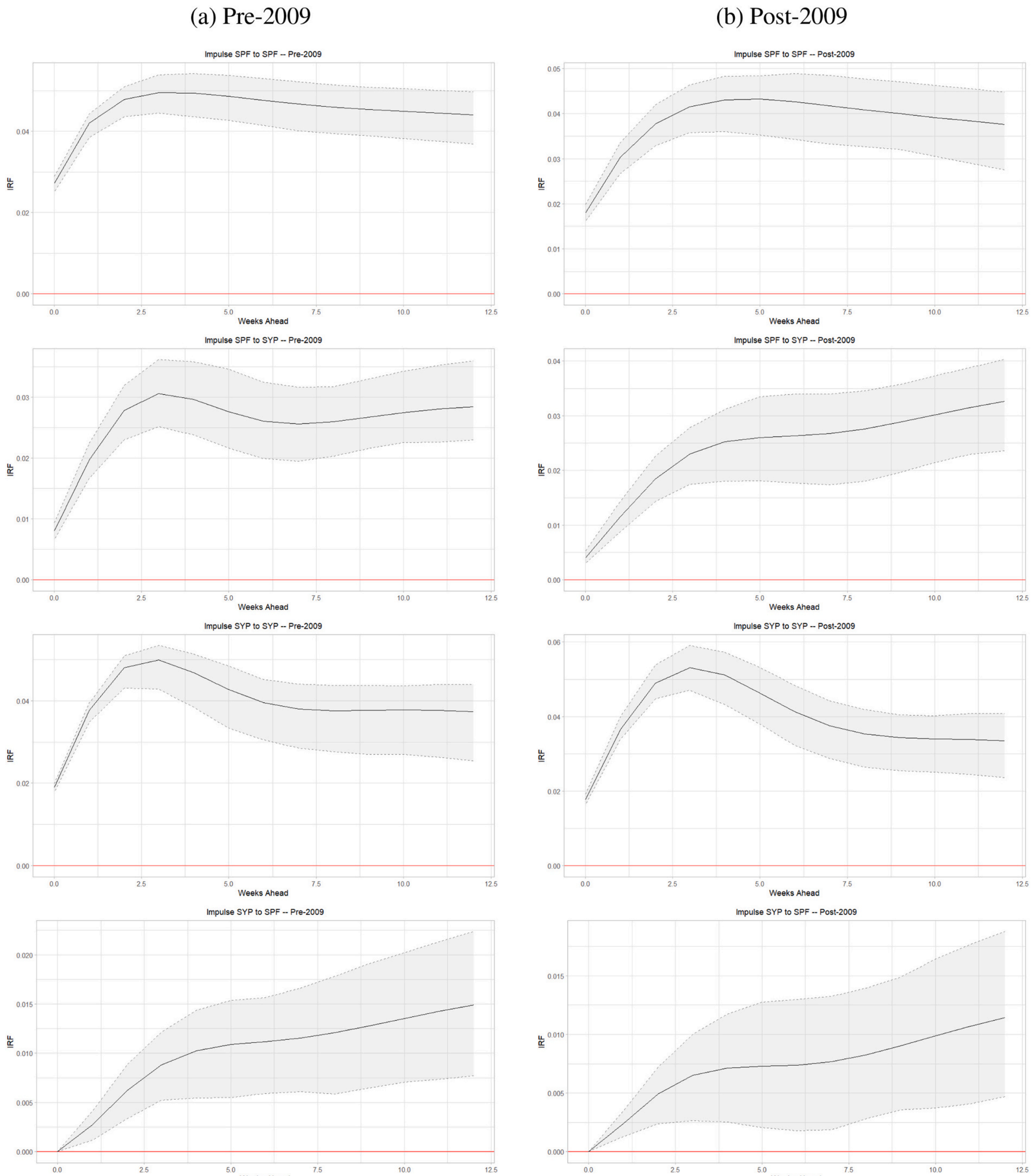

Fig. A.4. Orthogonalized Impulse Response of VEC Model – Pre-and Post-2009.
Note: This figure presents the regime-specific orthogonalized impulse response of VEC models using segmented data at 12/04/2009. Left panels under (a) present impulse responses before the break date, and right panels under (b) present impulse responses after the break date. IRFs are generated based on a VEC(2) model. The data frequency is weekly. All variables are transformed into natural logs. The shaded areas represent 95 % bootstrapped error band.

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