Essays on US Forest Products Markets

by

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Abstract

International trade of forest products plays a critical role in the economy of many countries around the world. Knowledge of market curves elasticities and the effect of currency exchange rates on goods trade are instrumental for gains, losses, and effective trade policies. Government cost-share programs have become important environmental policies around the world to reduce environmental degradation. This dissertation consists of three separate chapters on U.S. import of softwood plywood, the effect of exchange rate on U.S.-Brazil softwood plywood market, and the effect of the Conservation Reserve Program on U.S. timber market.

In the first chapter, we estimate U.S. import demand functions for softwood plywood from Brazil, Canada, Chile, and China using a cointegration approach. Both short run and long run U.S. import demand elasticities for softwood plywood under the evidence of cointegration are estimated. The results suggest that in the long-run, U.S. import demand for softwood plywood is more sensitive to own price change.

The second chapter investigates the effect of U.S. dollar's strengthening with respect to Brazil real on U.S.-Brazil softwood plywood market. An equilibrium displacement modelling approach is used. The price and quantity effects of U.S. dollar's strengthening is estimated using both a deterministic and stochastic approach. The results both in the short-run and long-run show that the pass-through effect of U.S. dollar's appreciation is more perceptible to U.S. consumers when U.S. faces a relatively elastic export supply curve from Brazil. Conversely, the absorption effect of U.S. dollar's strengthening is more perceptible to Brazilian producers when U.S. faces a relatively inelastic export supply curve from Brazil. Overall, the results show that U.S. dollar's strengthening with respect to Brazilian real during the period 2010-19 had a net positive welfare effect on U.S.-Brazil softwood plywood market.

The third chapter investigates the effect of the Conservation Reserve Program (CRP) on the equilibrium prices of standing timber in the U.S. South. For this purpose, a reduced form model for stumpage price is developed and estimated using data from 1960 to 2019. The time series econometrics that includes the tests of stationarity, the cointegration analysis and the linear model with stationary time series variables was used. The results of the estimated model suggest that CRP induced a significant increase in the supply of standing timber that depressed its equilibrium price after 24 years.

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List of Abbreviations

ADF: Augmented Dickey-Fuller

AEG: Augmented Engle and Granger

ARDL: Autoregressive Distributed Lag

BRL: Brazil Real

CND: Canadian Dollar

CRP: Conservation Reserve Program

ECM: Error Correction Model

FAS: Foreign Agriculture Service

FRED: Federal Reserve Bank Economic Data

KPSS: Kwiatkowski-Phillips-Schmidt-Shin

LM: Lagrange Multiplier

LR: Long run

OLS: Ordinary Least Square

PMG: Pool Mean Group

PPIACO: Producer Price Index for All Commodities

SBIC: Schwarz's Bayesian Information Criteria

SBP: Soil Bank Program

SR: Short run

USD: U.S. Dollar

USDA: U.S. Department of Agriculture

Chapter 1: U.S Import Demand for Softwood Plywood: A Cointegration Approach

Abstract

In this study, we present a model and estimates of U.S. import demand for softwood plywood from Brazil, Canada, Chile, and China using a cointegration framework. Our results suggest that the short-run own price elasticities of U.S. import demand from Brazil, Canada, Chile, and China are -0.53, -0.77, -0.45, and -0.90, respectively. Evidence of long-run cointegration are found for time series variables in U.S. import demand functions from Brazil, Canada, and China but not from Chile. The estimates of the long-run price elasticities suggest that U.S. import demands for softwood plywood from the three countries become more sensitive to softwood plywood price change.

Keywords: U.S. softwood plywood, import demand, time series order of integration, elasticity, substitute, complement.

Introduction

Softwood plywood is one of the most important forest products in the U.S. Invented by producers in U.S. Pacific Northwest in the early 1900s, its uses are mainly for residential and non-residential construction and industrial production (Mckeeve and Meyer 1983). Over the last century, the softwood plywood industry has expanded from the Pacific Northwest to the South and changed in many ways. The U.S. softwood plywood production and consumption peaked in the late 1980s, and as of 2017, they fell to about half of the peak level (Howard and Liang 2019). At the same time, substituting products such as oriented strand board are rising, and U.S. producers have faced competition of imports from other countries. For example, between 2010 and 2019, U.S. import for softwood plywood from the world increased by 346 percent (USDA-FAS 2020). Most of these imports are from Brazil, Canada, Chile, and China, which collectively accounted for 97% of total U.S. imports in 2017. In comparison, U.S imports for hardwood plywood and softwood lumber increased during the same period by only 7% and 55%, respectively (USDA-FAS 2020).

While the expansion of U.S. imports of softwood plywood is substantial, no estimates exist on the country's import demand for this product. In this study, we fill in this gap by estimating the U.S. import demand for softwood plywood from its major exporting countries. The estimates of import demand are pivotal for several reasons. First, they inform market actors and policymakers not only on the sensitivity of import demand and/or supply to key market factors but also on the potential effects of changes in these factors. Second, the knowledge of the relevant elasticity estimates is important for the assessment of the effectiveness of trade policies (Chou and Buongiorno 1982, Duffy et al. 1990). Third, the magnitudes of market elasticities provide important information on the distributive effects of shocks on the markets such as exchange rates (Zhang and Pearse 2011).

In the forest economics literature, several authors have estimated the import demand functions for several other forest products. For example, Chou and Buongiorno (1982) estimated the U.S. long-run aggregate import demand function for hardwood plywood. The results suggested that the import demand function was elastic with an estimated price elasticity at -1.98. The same authors in a second study (1983) estimated U.S. import long-run demand elasticities for hardwood plywood from Korea, Taiwan, Japan, and the Philippines at -1.95, -1.85, -1.81, and -1.77, respectively. Parthama and Vincent (1992) estimated the U.S. import demand elasticity of hardwood plywood from Indonesia at -2.81. Nagudabi and Zhang (2013) estimated U.S. import demand elasticity of softwood lumber from Canada at -0.81. Detomasi (1969) estimated the U.S. import demand elasticity for Canadian wood pulp at -12.05. These studies show that the import demand elasticities and functions vary with respect to the forest product and by country of origin.

The objective of this study is to estimate the U.S. import demand functions for softwood plywood from its main sources—namely Brazil, Canada, Chile, and China. The next section presents our theoretical model, followed by empirical models and data. The final sections present the estimation results and conclusions with practical implications of the study.

2. Theoretical model

Following Buongiorno et al. (1979), Nagudabi and Zhang (2013), Parajuli and Zhang (2016b), the U.S. import demand for softwood plywood from country j is defined as:

$$Q_{jt} = f(P_{jt},\ PUS_t, PRW_{jt}, HS_t, MR_t) \tag{1}$$

where Q_{jt} is the quantity of softwood plywood imported from country j which is Brazil, Canada, China, and Chile, respectively at time t; P_{jt} is the price of the imported softwood plywood from country j at time t; PUS_t is the U.S. domestic price of softwood plywood at time t; PRW_t is the price level of softwood plywood from the rest of the world on U.S. import market (excluding the import from country j under consideration) at time t; HS_t is the U.S. housing starts at time t; and MR_t is the U.S. mortgage rate at time t.

Equation (1) states that U.S. import demand for softwood plywood from country j is a function of the own prices of the imported plywood; the prices of U.S. domestic plywood and that of U.S. imported softwood plywood from the rest of the world, U.S. housing activities, and U.S. mortgage rates. All the import demand functions are assumed to be downward sloping. Hence, the coefficient for price of imported softwood plywood from country j (P_{jt}) is expected to be negative. The coefficient of the demand increasing factor—housing activity (HS_t) is expected to be positive. An increase in the mortgage rate increases the marginal cost of housing activities, and indirectly reduces the demand for softwood plywood for constructions. Hence, the coefficient of mortgage rate (MR_t) is expected to be negative. Finally, the coefficients of the prices of U.S. domestic softwood plywood (PUS_t), and the imports of softwood plywood from the rest of the world (PRW_t) may be positive or negative depending on whether the commodities are substitutes or complements to the imported softwood plywood from country j.

3. Empirical Models and Data

3.1 Empirical model

The log-log model is convenient for the estimation of the elasticities. Hence, Equation (1) can be rewritten as:

$$lnQ_{jt} = \alpha_{0j} + \alpha_{1j}lnP_{jt} + \alpha_{2j}lnPUS_t + \alpha_{3j}lnPRW_{jt} + \alpha_{4j}lnHS_t + \alpha_{5j}lnMR_t + u_{jt} \tag{2}$$

where α_0 is intercept; α_{1j} is the own price elasticity of U.S. import demand for softwood plywood from country j; $\alpha_{2j}\ldots\alpha_{5j}$ are the U.S. import demand elasticities from country j with respect to U.S. domestic softwood plywood price, the price level of U.S. import of softwood plywood from the rest of the world (excluding the country j under consideration), U.S. housing starts, and U.S. mortgage rate, respectively; and u is a random error term that includes all the other factors influencing U.S. import of plywood from country j.

Equation (2) ignores the dynamic effect of import demand factors such as past imports of softwood plywood or price expectations that may be important in forest products demand (Chou and Buongiorno 1982). To account for the dynamic effect of the import demand factors, Equation (2) can be rewritten as:

$$lnQ_{jt} = \alpha_{0j} + \sum_{i=1}^{k0} \varphi_{ji}\, lnQ_{j,t-i} + \sum_{i=0}^{k1} \alpha_{1ji}\, lnP_{j,t-i} + \sum_{i=0}^{k2} \alpha_{2ji}\, lnPUS_{t-i} + \sum_{i=0}^{k3} \alpha_{3ji}\, lnPRW_{jt-i} + \sum_{i=0}^{k4} \alpha_{4ji}\, lnHS_{t-i} + \sum_{i=0}^{k5} \alpha_{5ji}\, lnMR_{t-i} + \epsilon_{jt} \qquad (3)$$

where φ_{ji} is the coefficient for *i*th lagged import quantity from country *j*, and α_{1ji} is the *i*th lagged own price elasticity from country *j*, k0…k5 are lagged periods for various independent variables to be estimated; and all other parameters have a similar interpretation as in Equation (2).

To estimate Equation (3) for country j, the time series econometrics approach is used to avoid spurious regression. Thus, we first tested the stationarity of the time series variables using the Augmented Dickey Fuller (ADF) unit root test and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test. Second, given the order of integration of each time series variable in Equation (3) for country j, the cointegration of the time series is explored if applicable.

If the time series variables in Equation (3) are each integrated of order 1 for country j, the Augmented Engle and Granger (AEG) test for cointegration can be used to determine whether the variables are cointegrated (Engle and Granger 1987, Phillips and Ouliaris 1990). The AEG test of cointegration is a robust alternative to the multivariate Johansen tests for cointegration (MacKinnon 2010). Under the evidence of cointegration among the time series variables, Equation (3) can be rewritten in an error correction form as[1]:

$$\Delta lnQ_{jt} = a_{j0} + \lambda_j z_{j,t-1} + \sum_{i}^{s} \rho_{ji} \Delta lnQ_{j,t-i} + \sum_{i=0}^{k} \alpha_{1ji} \Delta lnP_{j,t-i}$$
$$+ \sum_{i=0}^{l} \alpha_{2ij} \Delta lnPUS_{t-i} + \sum_{i=0}^{m} \alpha_{3ij} \Delta lnPRW_{jt-i} + \sum_{i=0}^{n} \alpha_{4ij} \Delta lnHS_{t-i}$$
$$+ \sum_{i=0}^{r} \alpha_{5ij} \Delta lnMR_{t-i} + v_{jt} \tag{4}$$

where the parameters ρ_{ji}, α_{1ji}, α_{2ji}, α_{3ji}, α_{4ji}, α_{5ji} are the short-run parameters and λ_j, measuring the speed of adjustment of the U.S. softwood plywood import from country j. $z_{j,t-1}$ is the error correction term (ECM) of the import demand function for softwood plywood from country j. a and v are constant and error terms, respectively.

However, if the time series variables in Equation (3) for country j include both variables integrated of order 0 and 1, the autoregressive distributed lag (ARDL) bound test is recommended to explore the cointegration of the variables (Pesaran et al. 2001, Parajuli and Zhang 2016a). In this case, Equation (3) can be reparametrized in the following ARDL error correction form (Hassler and Wolters 2006):

$$\Delta lnQ_{jt} = a_{j0} + \sum_{i=1}^{s} \theta_{ji} \Delta lnQ_{ji,t-i} + \sum_{i=0}^{k} \delta_{1ji} \Delta lnP_{j,t-i} + \sum_{i=0}^{l} \delta_{2ji} \Delta lnPUS_{t-i}$$
$$+ \sum_{i=0}^{m} \delta_{3ji} \Delta lnPRW_{jt-i} + \sum_{i=0}^{n} \delta_{4ji} \Delta lnHS_{t-i} + \sum_{i=0}^{r} \delta_{5ji} \Delta ln\, MR_{t-i} + \omega_{0j} lnQ_{j,t-1}$$
$$+ \omega_{1j} lnP_{j,t-1} + \omega_{2j} lnPUS_{t-1} + \omega_{3j} lnPRW_{j,t-1} + \omega_{4j} lnHS_{t-1} + \omega_{5j} lnMR_{t-1} +$$
$$\varepsilon_{jt} \tag{5}$$

[1] Note that all dynamic models can be reparametrized in an error correction model (ECM) using Equations (4) or (5). However, the choice of the functional form of the ECM is recommended based on the order of integration of the time series variables in the model (Engle and Granger 1987, Pesaran and Shin 1999).

where θ, δ represent the short-run parameters in the model, and ω long-run parameters. α and ε are constant and error terms, respectively.

3.2 Data

Monthly data from 1995 to 2019 is used for model estimation in this study. This period is chosen because the monthly price data of U.S. domestic softwood plywood is available and because of a significant increase in the share of U.S. softwood plywood imports in U.S. consumption, from 1% in 1995 to nearly 17% in 2017 (Howard and Liang 2019). All the prices of softwood plywood are deflated to the base year 1995 using the Producer Price index for all commodities (Federal Reserve Economic Data 2020).

Table 1.1 presents the description and the sources of the data used in the study. Missing values were observed for price and quantity variables for softwood plywood imports from China and Chile and were thus removed for the import demand estimation for these countries.

Table 1.1. Variable description and sources of the data.

Variable	Description	Source
q_{br}	U.S. import of softwood plywood from Brazil (cubic meters)	USDA Foreign Agricultural Service Trade System (GATS) database (Accessed on April 15, 2020)
q_{cnd}	U.S. import of softwood plywood from Canada (cubic meters)	Same as above
q_{chl}	U.S. import of softwood plywood from Chile (cubic meters)	Same as above
q_{chn}	U.S. import of softwood plywood from China (cubic meters)	Same as above
PUS	Real price of U.S. domestic softwood plywood (USD/m^3) (=Weighted average of the quarterly price (USD/m^3) of U.S. southern pine and western fir plywood	Random lengths. 2005, 2009, 2015, 2019. Forest Product Market Prices and Statistics. 2005, 2009, 2015, 2019 Yearbook.
p_{br}	price (USD/m^3) of U.S. import of softwood plywood from Brazil (= Value of import /quantity of import from Brazil)	USDA Foreign Agricultural Service Trade System (GATS) database (Accessed on April 15, 2020)

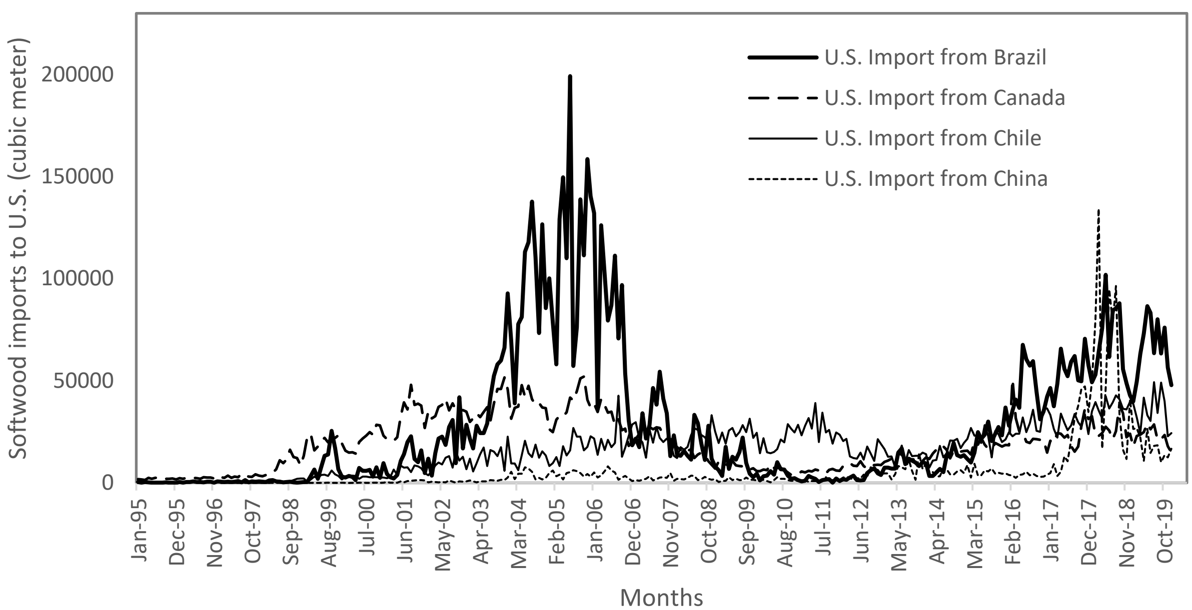
p_{cnd}	Real price (USD/m^3) of U.S. import of softwood plywood from Canada (= Value of import/quantity of import from Canada)	Same as above
p_{chl}	Real price (USD/m^3) of U.S. import of softwood plywood from Chile (= Value of import/ quantity of import from Chile)	Same as above
p_{chn}	Real price (USD/m^3) of U.S. import of softwood plywood from China (= Value of import/quantity of import from China)	Same as above
PRW_{br}	Price (USD/m^3) of U.S. import of softwood plywood from the rest of the world excluding imports from Brazil	Same as above
PRW_{cnd}	Price (USD/m^3) of U.S. import of softwood plywood from the rest of the world excluding imports from Canada	Same as above
PRW_{chl}	Price (USD/m^3) of U.S. import of softwood plywood from the rest of the world excluding imports from Chile	Same as above
PRW_{cnd}	Price (USD/m^3) of U.S. import of softwood plywood from the rest of the world excluding imports from Canada	Same as above
HS	U.S. housing start (new privately owned housing units started)	Federal Reserve economic data. https://fred.stlouisfed.org (accessed March 20, 2020)
MR	Real U.S. mortgage rate = ([1+ Nominal mortgage rate]/[1+inflation rate])-1 (Fisher 1977, Mankiw 2010).	Federal Reserve economic data. https://fred.stlouisfed.org (accessed March 20, 2020)

Table 1.2 presents a statistical summary of the data used. Figure 1.1 presents the historical trends of U.S. softwood plywood imports from Brazil, Canada, Chile and China.

Table 1.2. Summary statistics of variables.

Series	n	Mean	Std. deviation	Minimum	Maximum
q_{br}	299	28795.38	35894.20	0.00	199236.00
q_{cnd}	299	19231.94	12650.03	1411.00	52060.00
q_{chl}	287	16627.54	11581.10	0.00	49329.00
q_{chn}	275	6335.815	14999.42	0.00	134254.00
PUS	299	426.78	70.07	293.17	687.81
p_{br}	298	215.75	66.76	111.85	742.83
p_{cnd}	299	305.83	72.19	133.73	567.10
p_{chl}	263	294.62	68.04	181.32	907.24
p_{chn}	239	295.46	131.16	117.74	1060.79
PRW_{br}	299	296.52	55.91	185.70	532.11
PRW_{cnd}	299	256.93	60.29	152.37	555.45
PRW_{chl}	287	258.13	62.80	179.41	465.74
PRW_{chn}	275	256.99	49.13	156.03	517.87
HS	299	1314.73	433.83	478.00	2273.00
MR	299	5.66	1.49	3.35	9.12

Figure 1.1. Softwood plywood shipments to the U.S. from Brazil, Canada, Chile and China: 1995-2019.



Data source: USDA-FAS (2020).

4. Results and Discussion

4.1 Test of Stationarity

As noted earlier, the stationarity of each time series variable is examined using the KPSS and ADF unit root tests. The ADF unit root test is conducted for equations with intercept only, and intercept and linear trend. Table 1.3 presents the optimal lags number and the unit root test on the time series variables. The optimal number of lags is selected using Schwarz's Bayesian information criteria (SBIC).

The results in Table 1.3 suggest that all the time series variables are first difference stationary (integrated of order 1) except four variables: p_{cnd} (price of U.S. import of softwood plywood from Canada), q_{chl} (quantity of U.S. import of softwood plywood from Chile), p_{chl} (price of U.S. import of softwood plywood from Chile), and PRW_j (price of U.S. imports of softwood plywood from the rest of the world (excluding imports from country j = Brazil, Canada, Chile, and China). These variables are level stationary (integrated of order 0) due to the absence of unit roots.

We also used the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test to examine the stationarity of each time series variable. The maximum lag order is selected by combining the quadratic spectral kernel and automatic bandwidth selection (Hobijn et al., 1998). The KPSS differs from the ADF test by assuming the stationarity of the variables as null hypothesis. This test has been used in the economics literature to explore time series stationarity (Cheung and Chinn, 1997; Culver and Papell, 1999, Parajuli and Zhang, 2016). Table 1.4. presents the KPSS unit root test with intercept only, and intercept and linear trend. The results in Table 1.4 show strong evidence against the null hypothesis of stationarity for all the time series variables at 5%.

Table 1.3. Summary of the ADF unit root pretest on time series variables.

Variable	Specification	Lags	ADFc	ADFc,t
q_{br}	log	10	-2.012	-1.771
	first difference in log	9	-7.496[c]	-7.549[c]
q_{cnd}	log	1	-2.523	-2.217
	first difference in log	0	-20.793[c]	-20.857[c]
q_{chl}	log	10	-4.032[c]	-3.658[b]
	first difference in log	2	-13.060[c]	-13.225[c]
q_{chn}	log	5	-2.015	-2.527
	first difference in log	4	-9.954[c]	-9.943[c]
PUS	log	2	-2.905[b]	-3.007
	first difference in log	1	-13.162[c]	-13.143[c]
p_{br}	log	3	-1.747	-2.108
	first difference in log	2	-15.942[c]	-15.915[c]
p_{cnd}	log	1	-4.081[c]	-4.192[c]
	first difference in log	2	-14.117[c]	-14.141[c]
p_{chl}	log	1	-2.963[b]	-3.181[a]
	first difference in log	0	-19.375[c]	-19.340[c]
p_{chn}	log	4	-2.152	-2.010
	first difference in log	3	-11.545[c]	-11.557[c]
PRW_{br}	log	1	-3.78[c]	-4.06[c]
	first difference in log	0	-24.49[c]	-24.47[c]
PRW_{cnd}	log	1	-3.10[b]	-3.45[b]
	first difference in log	0	-24.96[c]	-24.92[c]
PRW_{chl}	log	1	-2.89[b]	-3.24[a]
	first difference in log	0	-25.01[c]	-24.97[c]
PRW_{chn}	log	1	-2.72[a]	-3.02
	first difference in log	0	-21.64[c]	-21.61[c]
HS	log	1	-1.076	-0.891
	first difference in log	0	-25.718[c]	-25.697[c]
MR	log	1	-1.539	-3.408[a]
	first difference in log	0	-13.274[c]	-13.255[c]

Notes: ADF_c and $ADF_{c,t}$ refer to ADF-t statistics when an intercept is included and when an intercept and time trend are included.

[a], [b], and [c] indicate that the null hypothesis of unit root is rejected at the 10%, 5% and 1% level, respectively.

Table 1.4. Summary of the KPSS unit root pretest on time series variables.

Variable	Specification	Lags	KPSSc	KPSSc,t
q_{br}	log	4	1.530[c]	0.643[c]
	first difference in log	4	0.086	0.063
q_{cnd}	log	4	0.986[c]	0.733[c]
	first difference in log	4	0.304	0.122[a]
q_{chl}	log	3	3.970[c]	0.869[c]
	first difference in log	3	0.102	0.026
q_{chn}	log	3	3.240[c]	0.424[c]
	first difference in log	3	0.025	0.026
PUS	log	4	1.190[c]	0.378[c]
	first difference in log	4	0.032	0.027
p_{br}	log	4	1.550[c]	0.672[c]
	first difference in log	4	0.056	0.057
p_{cnd}	log	4	1.620[c]	0.650[c]
	first difference in log	4	0.082	0.023
p_{chl}	log	3	0.775[c]	0.411[c]
	first difference in log	3	0.034	0.033
p_{chn}	log	3	1.100[c]	0.838[c]
	first difference in log	3	0.045	0.037
PRW_{br}	log	4	2.130[c]	0.654[c]
	first difference in log	4	0.052	0.032
PRW_{cnd}	log	4	1.960[c]	0.866[c]
	first difference in log	4	0.046	0.042
PRW_{chl}	log	3	0.537[b]	0.493[c]
	first difference in log	3	0.070	0.067
PRW_{chn}	log	4	1.300[c]	0.655[c]
	first difference in log	4	0.046	0.044
HS	log	4	2.100[c]	0.589[c]
	first difference in log	4	0.293	0.231[c]
MR	log	4	5.160[c]	0.260[c]
	first difference in log	4	0.036	0.028

Notes: $KPSS_c$ and $KPSS_{c,t}$ refer to KPSS test statistics when an intercept is included and when an intercept and time trend are included.
[a], [b], and [c] indicate that the null hypothesis of stationarity is rejected at the 10%, 5% and 1% level, respectively.

4.2 ARDL bound tests of cointegration

The results of the unit root test in Table 1.3 suggest that the time series variables in the U.S. import demand functions include both variables integrated of order 0 (p_{cnd}, q_{chl}, p_{chl}, PRW) and order 1 (q_{br}, p_{br}, q_{cnd}, p_{chn}, PUS, HS, MR). The ARDL bound test of cointegration is used to test whether the time series variables are cointegrated in their corresponding import demand function (Pesaran and Shin 1999, Pesaran et al. 2001). The ARDL bound test allows to have time series variables of different orders of integration. The bound test consists of estimating equation (5) using the Ordinary Least Square (OLS) method and testing two hypotheses (Hassler and Wolters 2006, Parajuli and Zhang 2016a). The first hypothesis test is an *F*-test with the null hypothesis defined as $\omega_0 = \omega_1 = \omega_2 = \omega_3 = \omega_4 = \omega_5 = 0$. The second hypothesis test is a *t*-test with the null hypothesis defined as $\omega_0 = 0$.

Table 1.5 presents the summary of the ARDL bound test of cointegration of the time series variables in the U.S. import demand functions. The optimal lag number is selected using SBIC. The Breusch-Godfrey LM test of serial correlation indicates that there is no evidence against the null hypothesis of independence of the residuals. Both the *F*- and *t*-statistics are statistically significant at 1% for the variables in the U.S. import demand function from Brazil and 5% for the variables in the U.S. import demand functions from Canada and China. This suggests that there is strong evidence of long-run cointegration between the time series variables in the U.S. import demand function from Brazil, Canada, and China. In contrast, for the variables in the U.S. import demand function from Chile, the test statics (*F*- and *t*-statistics) are not statistically significant at 5%. This suggests that there is no evidence of long-run cointegration between the time series variables in the U.S. import demand function from Chile.

Table 1.5. ARDL bounds test.

Model	ARDL (p, q_1, q_2, q_3, q_4, q_5)	p-value, LM test [a]	F-statistics	t-statistics	*Cointegration?*
Import demand from Brazil [d]	*(2,1,1,1,1,1)*	*0.65*	12.640[c]	-7.962[c]	Yes
Import demand from Canada [d]	*(3,1,1,1,1,1)*	0.29	5.500[c]	-4.635[b]	Yes
Import demand from Chile [e]	*(4,1,1,1,1,1)*	0.12	3.574	-3.920	No
Import demand from China [f]	*(3,1,1,1,1,1)*	0.25	4.435[b]	-4.839[c]	Yes

[a] Breusch-Godfrey serial correlation LM test on the residuals at lags 1, 3, 4, 3 for U.S. import function from Brazil, Canada, Chile and China, respectively.
[b] rejection of the null hypothesis at 5% level of significance. [c] rejection of the null hypothesis at 1% level of significance.
[d] The lower and upper bound values for the *F*-statistics at the 1 and 5% level of significance are 3.44 and 4.77, and 2.64 and 3.82, respectively (Pesaran et al., 2001)
[e] The lower and upper bound values for the *F*-statistics at the 1 and 5% level of significance are 3.44 and 4.79, and 2.64 and 3.83, respectively (Pesaran et al., 2001)
[f] The lower and upper bound values for the *F*-statistics at the 1 and 5% level of significance are 3.46 and 4.81, and 2.65 and 3.84, respectively (Pesaran et al., 2001).

4.3. U.S. Import Demand from Brazil, Canada and China

Table 1.6 presents the long run and short run elasticities' estimates of U.S. import demand from Brazil, Canada, and China using the ARDL Error Correction model (ECM). The ECM coefficients (adjustment parameters) in Table 1.6 are negative and statistically significant at 1%, suggesting satisfactory fits of the ECM models. Furthermore, the elasticities in the import demand functions have the expected signs.

The short-run own price elasticities of the U.S. import demand for softwood plywood from Brazil, Canada, and China are negative as expected and statistically significant at 1%. This suggests that U.S. import demand for softwood plywood responds to changes in the price of softwood plywood from Brazil, Canada, and China within a month following the price change. Particularly, a 1% increase in the price of Brazilian softwood plywood is associated with a 0.53% decrease in U.S. import demand for plywood from Brazil. Similarly, a 1% increase in the price of softwood plywood from Canada is associated with a 0.77% decrease in U.S. import demand of plywood from Canada. For U.S. import demand of softwood plywood from China, a 1% increase in own price is associated with 0.90 decreases in the U.S. import demand. This suggest that in the short-run, U.S. import demand for softwood plywood from Brazil is relatively inelastic compared to U.S. import demand from Canada and China.

The short-run cross-price elasticity of U.S. import demand from Brazil with respect to the price of U.S. domestic softwood plywood is negative and statistically significant at 1%. This suggests that in the short-run, softwood plywood from U.S. and Brazil are complement. A 1% increase in the price of U.S. domestic softwood plywood is associated with a 3.24% decrease in U.S. import of Brazilian plywood. In contrast, for U.S. import demand from Canada, the short-run cross price elasticity with respect to U.S. domestic plywood is positive. This suggest that in the short-run, U.S. and Softwood plywood from Canada are close substitutes. A 1% increase in the price of U.S. domestic plywood is associated with a 0.40% increase in U.S. import of Canadian softwood plywood. No evidence is observed on the response of U.S. import demand for softwood plywood from China with respect to changes in the price of U.S. domestic production in the short-run. Furthermore, no evidence is observed on the response of U.S. import demand for softwood plywood from Brazil, Canada, and China with respect to changes in the price of U.S. imports from the rest of the world, U.S. housing conditions within a month following changes in these variables.

The long-run own price elasticities of U.S. import demand for softwood plywood from Brazil, Canada, and China are negative as expected and statistically significant at 1% and 5%, respectively. Furthermore, the long-run own price elasticities of the U.S. import demand functions from the three countries are more than twice their corresponding short-run elasticities.

This indicates that the U.S. import demand functions are more elastic in the long-run relative to the short-run U.S. import demand functions. This is consistent with economic theory which suggests that consumers and producers are more responsive to changes in commodity prices when they have more time to respond (Varian 1992, Zhang and Pearse 2011, Kinnucan 2016). Particularly, a 1% increase in the own-price of softwood plywood from Brazil, Canada, and China is associated with a decrease of 4.04, 2.74, and 2.67 of U.S. import from these countries, respectively.

The long-run cross price elasticities of U.S. import demand from Brazil, China, and Canada with respect to U.S. domestic softwood plywood prices are positive and statistically significant at 1% and 5%, respectively. This suggests that it takes several months to U.S. demand for softwood plywood to substitute U.S. domestic production with softwood plywood imports from Brazil, Canada, and China following an increase in the prices of the U.S. domestic softwood plywood. Particularly, a 1% increase in the price of U.S. domestic softwood plywood is associated in the long-run with 7.66%, 2.28%, and 6.66% increase of U.S. import from Brazil, Canada, and China, respectively.

The long-run cross price elasticities of U.S. import demand from Brazil and China with respect to the price level of the rest of the world softwood plywood on U.S. import market are negative and statistically significant at 1% and 10%, respectively. This suggests that softwood plywood from Brazil, China, and the rest of the world are complements on U.S. softwood plywood import market in the long-run. The long-run elasticities of U.S. import demand from Brazil, Canada, and China with respect to U.S. mortgage rates are negative as expected and statistically significant at 1%. This suggests that the cost increasing effect of mortgage rate on U.S. housing market becomes perceptible several months after an increase in mortgage rates.

Similarly, the effects of changes in U.S. housing starts are perceptible on U.S. softwood plywood import market several months after changes in the housing conditions. Particularly, an increase in U.S. housing starts is associated with an increase of U.S. import of softwood plywood from Brazil and Canada as opposed to China. This may suggest that softwood plywood from Brazil and Canada are close substitutes to U.S. domestic production than softwood plywood from China in the long-run.

Table 1.6. ARDL ECM model estimates of the elasticities of U.S. import demand from Brazil, Canada and China.

	Import demand function					
	Brazil		Canada		China	
ADJ parameter	-0.276[c]	(0.037)	-0.105[c]	(0.025)	-0.190[c]	(0.041)
Long-run estimates						
$\ln p_j$	-4.040[c]	(0.551)	-2.747[c]	(0.544)	-2.665[b]	(1.245)
$\ln PUS$	7.661[c]	(1.233)	2.280[c]	(0.873)	6.659[c]	(2.010)
$\ln PRW_j$	-4.516[c]	(0.975)	-0.108	(0.684)	-4.159[b]	(1.942)
$\ln HS$	0.959[a]	(0.487)	1.558[c]	(0.407)	-2.390[a]	(1.216)
$\ln MR$	-2.430[c]	(0.517)	-1.137[c]	(0.417)	-3.192[c]	(1.011)
Short-run estimates						
Δlnp_j	-0.532[b]	(0.227)	-0.773[c]	(0.088)	-0.900[c]	(0.222)
$\Delta lnPUS$	-3.244[c]	(0.716)	0.395[a]	(0.201)	0.822	(0.791)
$\Delta lnPRW_j$	0.421	(0.395)	-0.065	(0.094)	-0.363	(0.526)
$\Delta lnHS$	-0.420	(0.400)	-0.131	(0.119)	-0.456	(0.432)
$\Delta lnMR$	2.420[a]	(0.889)	0.055	(0.263)	1.212	(1.010)

[a] Indicates significance levels at the 10% level of probability. [b] Indicates significance levels at the 5% level probability. [c] Indicates significance levels at 1% level probability.
j= Brazil, Canada, and China
Standard errors in parentheses

4.4. U.S. Import Demand Function from Chile

Table 1.7 presents the U.S. import demand function for softwood plywood from Chile. Because the ARDL bound test in Table 1.5 indicates no evidence of cointegration relationship among the time series variables in the import demand function from Chile, the estimated U.S. import demand function represents a short-run function only. The variables are included in the import demand equation following their order of integration in Table 1.3 to avoid a spurious regression. The lag variable in the import demand equation is included to remove the serial correlation in the error term.

Table 1.7. U.S. import demand function from Chile

	Dependent variable: $\ln q_{chile}$	
Variables	Coefficient	S.E
$\ln p_{chl,t}$	-0.446[b]	(0.213)
$\ln PUS$	0.123	(0.188)
$\ln PRW_t$	-0.224[a]	(0.186)
$\Delta lnHS$	0.041	(0.301)
$\Delta lnMR$	0.591	(0.737)
$\ln q_{chl,t-1}$	0.871[c]	(0.032)
constant	4.215[c]	(1.552)
Feb.	-0.074	(0.117)
Mar.	0.058	(0.118)
Apr.	0.083	(0.117)
May	0.117	(0.116)
Jun.	0.008	(0.116)
Jul.	0.095	(0.115)
Aug.	0.038	(0.114)
Sep.	0.054	(0.115)
Oct.	0.015	(0.115)
Nov.	0.035	(0.116)
Dec.	0.012	(0.117)
Adj R-squared	0.88	

[a] Indicates significance levels at 10% level of probability
[b] Indicates significance levels at 5% level of probability
[c] Indicates significance levels at 1% level of probability

The results in Table 1.7 indicate that the U.S. import demand function for softwood plywood from Chile is downward sloping as expected with a coefficient significant at 5%. Furthermore, the softwood plywood from Chile is a substitute for softwood plywood from the rest of the world on U.S. import market. All other coefficients have the expected signs with the exception of that for mortgage rate, which is statistically insignificant.

4.5 Comparison

4.5.1. Comparison with Chou and Buongiorno (1982)

For comparison, the price elasticities of U.S. import demand functions are also estimated using a dynamic model proposed by Chou and Buongiorno (1982). Table 1.8 compares our results with the price elasticities following Chou and Buongiorno (1982) dynamic model. Model 1 shows the price elasticities we obtained using the cointegration approach in Tables 1.6 and 1.7 of this study. Model 2 shows the corresponding price elasticities with no monthly dummies in Chou and Buongiorno (1982) dynamic model using our data. In Model 3, price elasticities are estimated using the same regression as Model 2 with the monthly dummies included to account for monthly effect. Table 1.8. Long-run (LR) and Short-run (SR) price elasticities comparisons using ARDL cointegration and the dynamic model in Chou and Buongiorno (1982)

Model	Models	U.S. import demand from			
		Brazil	Canada	Chile	China
Model 1	Price elasticity SR	-0.532[b]	-0.773[c]	-0.446[b]	-0.900[c]
	Price elasticity LR	-4.040[c]	-2.747[c]	—	-2.665[b]
Model 2	Price elasticity SR[d]	-1.358[c]	-0.594[c]	-0.574[c]	-0.612[c]
	Price elasticity LR[d]	-4.456[c]	-2.827[c]	-2.481[c]	-2.079[c]
Model 3	Price elasticity SR[e]	-1.306[c]	-0.589[c]	-0.561[c]	-0.706[c]
	Price elasticity LR[e]	-4.419[c]	-2.948[c]	-2.458[c]	-2.348[c]

[a] Indicates significance levels at 10% level of probability. [b] Indicates significance levels at 5% level of probability. [c] Indicates significance levels at 1% level of probability.
[d] No monthly dummies included in Chou and Buongiorno (1982) dynamic model
[e] Monthly dummies included in Chou and Buongiorno (1982) dynamic model

The results in Table 1.8 show that the sign of the price elasticities of the U.S. import demand functions are consistent in the three models. However, while the long-run price elasticities are rather similar among all three model, the magnitudes of the short-run price elasticities vary quite significantly. These observed differences in the magnitudes of the price elasticities between our approach and Chou and Buongiorno (1982) dynamic model could be explained by the imposition of the same functional form to all the time series data ignoring the order of integration of the time series variables in Chou and Buongiorno (1982) approach. We believe that our results are more reliable than the dynamic model.

Interestingly, our estimates for short-run U.S. import demand elasticities for softwood plywood from Brazil (-0.53), Canada (-0.77), Chile (-0.45), and China (-0.90) are comparable to the estimates of -0.81 and -0.74 in U.S. import demand for softwood lumber from Canada in Nagubadi and Zhang (2013) and Parajuli and Zhang (2016b), respectively. This may suggest that softwood lumber and softwood plywood respond to similar market dynamics and conditions. Also, our long-run elasticity estimates for Brazil, Canada and China are comparable to the hardwood import demand elasticity estimates found by Chou and Buongiorno (1982, 1983).

4.5.2. Panel cointegration approach

For comparison, the panel cointegration approach is also used to estimate the U.S. import demand functions for softwood plywood. Under panel data framework, variables cointegration can be explored using panel cointegration tests (Harb 2005, Im and al. 2003, Pedroni 1999, Kao and al. 1999). To this end, the stationarity of the variables is explored using the panel unit root test approach. The Im-Pesaran-Shin (IPS) unit root test is used to explore the unit root of the variables as it allows heterogeneity among panel units and unbalanced panel (Im and al. 2003, Apergis and Payne 2010).

Table 1.9 reports the panel unit root test results, which suggest that all the variables are level stationary except the variables housing starts (HS) and mortgage rate (MR). These variables (HS, MR) are first difference stationary. Given that none of these variables is I(2), Pedroni's heterogenous panel tests of cointegration are used to explore the variable cointegration (Pedroni 1999, 2001, 2004). These tests consist of two sets of panel cointegration tests. The first set of tests are within dimension approach that includes four statistics: panel v, panel ρ, panel t, and panel ADF statistics. These statistics are derived from pooling autoregressive statistics across the countries and taking into account common time factors and heterogeneity across countries (Pedroni 2004, Apergis and Payne 2012). The second set of tests are between dimension approach that includes three statistics: group ρ, group t, and group ADF statistics. These statistics are derived from the averages of the individual autoregressive coefficients of the unit root of the residuals of each country in the panel (Pedroni 2004, Apergis and Payne 2012). Both set of statistics are distributed asymptotically as standard normal (Pedroni 1999).

Table 1.9. Panel unit root tests

Variable	Specification	IPS_c	$IPS_{c,t}$
qm	log	-2.92^b	-2.60^b
	First difference in log	-42.45^b	-45.50^b
pm	log	-3.11^b	-2.59^b
	First difference in log	-39.84^b	-42.47^b
PUS	log	-4.60^b	-3.88^b
	First difference in log	-26.95^b	-28.19^b
PRW	log	-3.93^c	-3.52^c
	First difference in log	-55.65^c	-60.01^c
HS	log	1.05	3.31
	First difference in log	-34.62^b	-36.72^b
MR	log	-0.02	-3.16^b
	First difference in log	-27.39^b	-28.67^b

IPS_c and $IPS_{c,t}$ refers to the Im-Pesaran-Shin test when an intercept and when intercept and time trend are included, respectively.
[b] indicates that the null hypothesis of unit root is rejected at 1% level.
qm = softwood plywood import quantity, pm = price of U.S. import for softwood plywood.

Table 1.10 reports both the panel and group mean cointegration test statistics. Five statistics out of the seven reject the null hypothesis of no cointegration at the 5% significance level. This suggests that there is evidence of long-run cointegration relationship among the variables. Given that the variables in the U.S. import demand function include both I(0) and I(1) variables, the panel ARDL model can be used to estimate the import demand function (Pesaran and al. 1997). Several estimators are generally used for the estimation of the panel ARDL model: the mean group estimator (MG) and the pool mean group estimator (PMG). While both estimators constrain all the long-run parameters across the panel units to be the same, the PMG estimator allows the short-run parameters and error variance to differ across panel units (Pesaran and al. 1999). In this study, the pool mean group estimator is selected as most appropriate estimator of the panel ARDL model using the Hausman test (Hausman 1978).

Table 1.10. Panel cointegration tests

Panel test statistics		Group mean panel test statistics	
Panel v-statistic	0.54		
Panel ρ-statistic	-5.97[c]	Group ρ-statistic	-7.07[c]
Panel t-statistic	-5.64[c]	Group t-statistic	-6.59[c]
Panel ADF statistic	1.54	Group ADF statistic	2.15[b]

[a] Indicates rejection of the null hypothesis of no cointegration at 10% significance level.
[b] Indicates rejection of the null hypothesis of no cointegration at 5% significance level.
[c] Indicates rejection of the null hypothesis of no cointegration at 1% significance level.

Table 1.11 presents the elasticities of U.S. import demand functions using the panel ARDL approach. The results indicate that the U.S. long run import demand function is downward sloping as expected and consistent with the time series approach results. Similarly, all the short-run own price elasticities are negative with the exception of that of U.S. import demand from Chile which is not statistically significant at 5%. Furthermore, the long-run own price elasticity of U.S. import demand is larger in magnitude relative to the short-run own price elasticities of U.S. imports from Brazil, Canada, and China. This result is consistent with the long-run elasticities estimated using the time series approach and confirms that the U.S. consumers and producers in the softwood plywood import market becomes more responsive to the change on softwood plywood prices when they have more time to respond. However, because the ARDL model estimated using the PMG estimator assumes homogeneity in the long run elasticities across the countries, only one long-run import demand function is estimated. Note that the long-run elasticities have the same sign and magnitude as the U.S. long-run elasticities from Brazil in the country-specific time series approach. Finally, though the panel ARDL estimates are relatively consistent with the estimates from the time series approach, estimates from the latter approach seem to be more reliable as the approach imposes no restrictions on the import demand functions parameters. For example, the long-run homogeneity assumption of the panel ARDL model could be strong as this suggests that consumers on U.S. import market consider softwood plywood from the four countries as identical.

Table 1.11. Long-run (LR) and Short-run (SR) elasticities of U.S import demand functions using panel ARDL approach.

Variables	LR elasticities	SR elasticities			
		Brazil	Canada	Chile	China
pm	-4.516[c]	-0.427[c]	-0.899[c]	0.370	-0.990[c]
PUS	7.590[c]	-2.922[c]	0.695[c]	-0.763[a]	-0.463
PRW	-3.037[c]	0.375	-0.069	0.124	-0.567
HS	0.586	-0.425	0.004	-0.243	-0.504
MR	-2.178[c]	2.563[c]	0.088	0.530	0.371

[a] Indicates significance levels at 10% level of probability. [b] Indicates significance levels at 5% level of probability. [c] Indicates significance levels at 1% level of probability.
pm = Price of U.S. import for softwood plywood.

Conclusions

In this study we estimate the U.S. import demand functions for softwood plywood from Brazil, Canada, Chile, and China which collectively account for more than 90 percent of U.S. imports in the 2010s. To this end, a cointegration framework is adopted following the order of integration of the time series variables in the import demand functions. As evidence for cointegration was found for the time series variables in the U.S. import demand functions for softwood plywood from Brazil, Canada, and China, both short-run and long-run elasticities are estimated for U.S. imports demand from these countries. However, only short-run U.S. import demand function is estimated for softwood plywood imports from Chile.

Our results suggest that the short-run own price elasticities of U.S. import demand functions from Brazil, Canada, Chile, and China are -0.53, -0.77, -0.45 and -0.90, respectively. In our model framework, the short-run elasticities show the responses of the U.S. import demand functions to changes in the model variables within the month following the price change.

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Chapter 2: Currency Exchange Rate Effect on US-Brazil Softwood Plywood Markets

Abstract

In this study, I investigate the effect of currency exchange rate on U.S.-Brazil softwood plywood market and the welfare distribution of U.S. dollar's strengthening with respect to Brazilian real between 2010 and 2019. The Equilibrium displacement modelling approach is adopted. The price and quantity effects of U.S. dollar's strengthening is estimated using both a deterministic and stochastic approach. The results both in the short-run and long-run show that the pass-through effect of U.S. dollar's appreciation is more perceptible to U.S. consumers when U.S. faces a relatively elastic export supply curve from Brazil. Conversely, the absorption effect of U.S. dollar's strengthening is more perceptible to Brazilian producers when U.S. faces a relatively inelastic export supply curve from Brazil. U.S. import demand for Brazilian plywood is larger in the long-run while the fall in U.S. domestic production of plywood is smaller in the short-run. Although, the U.S. dollar's strengthening with respect to Brazilian real had a detrimental effect on U.S. domestic producers of softwood plywood, the gains to U.S. consumers more than offset the losses to U.S. producers. Using the results from the stochastic approach, the net welfare gain to U.S. on the softwood plywood market is estimated at US $ 0.14 billion and US $ 1.21 billion in the short-run and long-run, respectively. The gain to Brazilian producers of U.S. dollar's strengthening is estimated at US $ 0.23 billion and US $ 1.66 billion in the short-run and long-run, respectively.

Keywords: Softwood plywood, exchange rate, pass-through elasticity, absorption, welfare analysis.

Introduction

Trade of forest products is a key component in the economy of many countries. One important factor that may influence trade between countries is the exchange rate that allows the conversion of foreign currency into domestic currency. In many parts of the world, not only currencies in exporting countries and importing countries are different, but they are also not static (Evans, 2011). Hence, understanding the effect of exchange rate on goods trade is of great importance for policymakers.

Between the U.S. and its major forest products trade partners, the largest appreciation of U.S. dollar (USD) during the last decade is observed with respect to Brazilian real (BRL). For example, between 2010 and 2019, the USD appreciated with respect to BRL by 124% (Figure 2.1). With respect to Canadian dollar (CND), the USD appreciated by 29% between 2010 and 2019 while with respect to Chinese Yuan (CNY), the USD appreciated by only 2%. In the same time, U.S. imports of softwood plywood from Brazil increased by 1796% (200% annually), one of the fastest among U.S. imports of forest products (Figure 2.1).

According to Schuh (1974), an overvalued currency acts as an implicit export tax for exporting countries and an implicit subsidy for importing countries. In this study, we test this hypothesis on a forest product market. Of particular interest is the extent to which the USD appreciation contributed to the observed rise of U.S. softwood plywood imports from Brazil during the period 2010-19 (Figure 2.1). Second, we investigate the welfare distribution of this implicit subsidy on U.S. Brazil softwood plywood trade markets.

Investigating the relationship between exchange rate and U.S. imports of softwood plywood is important for several reasons. First, softwood plywood is one of the most important forest products for U.S., and imports from Brazil is the largest source of competition for U.S.

domestic plywood (USDA FAS, 2020). Second, while USD' appreciation with respect to BRL is expected to hurt U.S. domestic production, these negative effects can occur if changes in the exchange rates are reflected in the price (in USD) of U.S. imports from Brazil. With the exception of Yang (1997) who estimated the aggregated exchange rate pass through on U.S. hardwood veener and plywood market, no study investigated pass through on U.S. softwood plywood market. Third, the exchange rate pass through has major trade policy implications (Uusivuori and Buongiorno, 1991; Hanninen, 1999; Marazzi and Sheets, 2007). For example, if USD has a negligeable pass through on softwood plywood market, currency realignments should only have a modest effect on imports and/or exports.

In the forest economics literature several authors investigated the relationship between changes in currency exchange rates, imports, and exports of forest products. For example, Uusivuori and Buongiorno (1990) investigated the short-run and long-run effects of changes in exchange rates on U.S. imports of papers from Finland and Sweden. Buongiorno et al. (1988) and Bolkesjo and Buongiorno (2006) analyzed the effects of exchange rates on U.S. imports of softwood lumber from Canada. Other studies include Kim et al. (2003) and Sun and Zhang (2003). The main conclusion from these studies is that the effects of exchange rates on commodities trade are ambiguous, commodity specific, and vary with respect to the time horizon. Changes in exchange rates generate both uncertainty and opportunities for profit when combined may reduce or increase trade between countries (Zhang and Kinnucan, 2014; Bahmani-Oskooee and Hegerty, 2014).

The extent to which changes in nominal exchange rates is reflected in the forest products price in destination countries' currencies also attracted the attention of some researchers. For example, Uusivuori and Buongiorno (1991) estimated the pass through of exchange rates on the prices of U.S. forest product exports to Europe and Japan. Similarly, Hanninen (1998) estimated the pass through of exchange rates on the prices of Finnish sawnwood in the UK market. Goodwin et al. (2012) investigated the pass through of USD appreciation and depreciation on the price of U.S. imports of oriented strand board (OSB) from Canada. The main conclusion of this strand of literature is that first, the exchange rates pass through is industry specific and bounded between zero and one. Second, exchange rate pass through is influenced by the economy stage of countries, and whether single or aggregate commodities are considered (Goodwin et al., 2012; Bussiere, 2013).

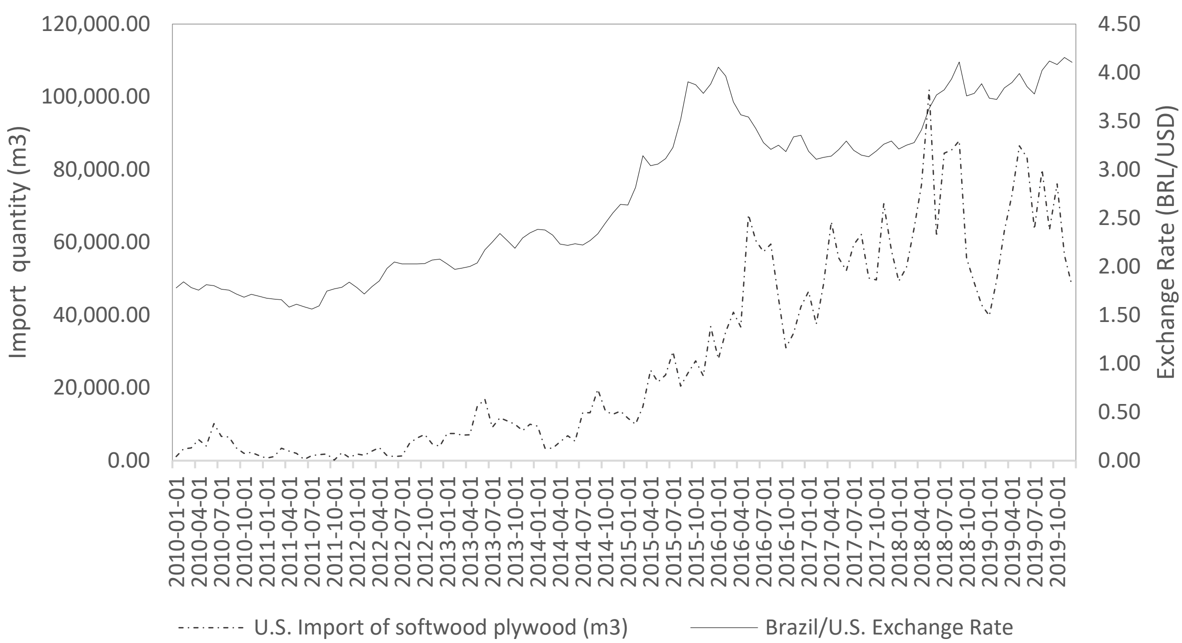

Figure 2.1. U.S./Brazil exchange rate and U.S. import of Brazilian softwood plywood, 2010 to 2019[2]

[2] U.S. import data is obtained USDA Foreign Agricultural Service Global Agricultural Trade System (GATS) database (accessed on May 22, 2021). Exchange rate data is obtained from the Federal Reserve Bank of St. Louis, Economic data, https://fred.stlouisfed.rog (accessed on May 22, 2021).

2. Theoretical model and Welfare formulas

2.1. Theoretical Model

To gain an insight on the effects of changes in exchange rate on U.S.-Brazil softwood plywood market, a set of demand and supply equations describing softwood plywood market is defined following Menon (1995). The model assumes that Brazilian and U.S. softwood plywood are perfect substitutes for each other. Furthermore, softwood plywood market is assumed competitive. U.S. is defined as a net importing country of softwood plywood and Brazil, a net exporting country. Because this study focuses on the effect of U.S. dollar's strengthening on U.S. imports of softwood plywood from Brazil, only imports from Brazil are considered. With the above assumptions, U.S. softwood plywood market can be described using the following set of structural equations:

$Q_d = D(P_{us})$ U.S. demand for softwood plywood (1)

$Q_s = S(P_{us})$ U.S. production for softwood plywood (2)

$Q_m = M(P_{br})$ U.S. excess demand for Brazil softwood plywood (3)

$P_{br} = P_{us}. e$ (4)

$Q_d = Q_s + Q_m$ (Market clearing condition) (5)

where P_{us}, and P_{br} are the prices of softwood plywood on U.S. market and that received by Brazilian producers, respectively; e ($e = \frac{BRL}{USD}$) denotes the exchange rate between Brazilian real and U.S. dollar. An increase in e implies that the U.S. dollar has strengthened with respect to the Brazilian real, which makes Brazilian softwood plywood less expensive for U.S. buyers; conversely, a decrease in e implies that the U.S. dollar has weakened with respect to Brazilian real. Q_d, Q_s, and Q_m represent U.S demand for softwood plywood, U.S. domestic production for softwood plywood, and U.S. import for softwood plywood, respectively.

The model contains five endogenous variables (Q_d, Q_s, Q_m, P_{us}, P_{br}) and one exogenous variable, e. Other exogenous variables that affect softwood plywood market curves are suppressed. To derive the effect of changes in exchange rate the Muth-type approach is used. We first express the model in proportionate change form by taking the total differential of each equation to yield:

$$Q_d^* = \eta_{us} P_{us}^* \tag{6}$$

$$Q_s^* = \varepsilon_{us} P_{us}^* \tag{7}$$

$$Q_m^* = \varepsilon_m P_{br}^* \tag{8}$$

$$P_{br}^* = P_{us}^* + e^* \tag{9}$$

$$Q_d^* = k_s Q_s^* + k_m Q_m^* \tag{10}$$

where $Z^* = dZ/Z$. In the above equations, η_{us} (< 0) is the U.S. demand elasticity for softwood plywood; $\varepsilon_i (> 0)$ $i = U.S., m$ are the supply elasticities for softwood plywood corresponding to each source (U.S., and Brazil); $k_s = Q_s/Q_d$ is the share of U.S. softwood plywood consumption from domestic source, and $k_m = Q_m/Q_d$ is the share of U.S. softwood plywood consumption from Brazilian source. The quantity shares sum to one $(k_s + k_m = 1)$.

Next, we solve for the U.S. import demand curve (Q_m^*) using equations (6), (7), and (10) to obtain:

$$Q_m^* = \left(\frac{\eta_{us} - k_s \varepsilon_{us}}{k_m}\right) P_{us}^* \tag{11}$$

Equation (11) can be rewritten by combining equations (9) and (11)

$$Q_m^* = \eta_m P_{br}^* - \eta_m e^* \tag{12}$$

with $\eta_m = \frac{\eta_{us} - k_s \varepsilon_{us}}{k_m} < 0$. Equation (12) states that the U.S. import demand for Brazilian softwood plywood is downward sloping as expected. Furthermore, Equation (12) suggests that holding the price of Brazilian softwood plywood constant, a strengthening of U.S. dollar with respect to Brazilian real increases U.S. import for Brazilian softwood plywood.

Equating equations (8) and (12) yield the reduced form equation for the price of Brazilian softwood plywood:

$$P_{br}^{*} = \frac{\eta_m}{\eta_m - \varepsilon_m} e^{*} \tag{13}$$

Equation (13) states that a strengthening of U.S. dollar increases the price received by Brazilian softwood plywood producers, but by less than the amount of the strengthening. Furthermore, the absorption elasticity to Brazilian producers of U.S. dollar strengthening is bounded between 0 and 1 [3].

The reduced form for the price of softwood plywood paid by U.S. consumers for Brazilian plywood is obtained by combining equations (9) and (13):

$$P_{us}^{*} = \frac{\varepsilon_m}{\eta_m - \varepsilon_m} e^{*} \tag{14}$$

Equation (14) states that a strengthening of U.S. dollar with respect to Brazilian real decreases the price paid by U.S. consumers for Brazilian plywood. Furthermore, the pass-through elasticity of U.S. dollar's strengthening to U.S. domestic consumer is bounded between -1 and 0.[4]

The reduced form equation for the quantity for U.S. import of Brazilian softwood plywood is obtained by substituting Equation (13) into Equation (8):

$$Q_m^{*} = \frac{\varepsilon_m \eta_m}{\eta_m - \varepsilon_m} e^{*} \tag{15}$$

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[3] This bound can be seen by dividing Equation (13) through by η_m: AE = $\frac{1}{1 - \frac{\varepsilon_m}{\eta_m}}$ where the upper limit holds when $\varepsilon_m = 0$ and the lower limit as $\varepsilon_m \to \infty$.

[4] This bound can be seen by dividing Equation (14) by ε_m: PTE = $\frac{1}{\frac{\eta_m}{\varepsilon_m} - 1}$ where the lower limit holds when $\eta_m = 0$ and the upper as $\eta_m \to \infty$.

Equation (15) states that a strengthening of U.S. dollar's with respect to Brazilian real increases U.S. import for Brazilian softwood plywood.

Similarly, the reduced form equation for U.S. domestic production for softwood plywood is obtained by substituting Equation (14) into Equation (7):

$$Q_s^* = \frac{\varepsilon_{us}\varepsilon_m}{\eta_m - \varepsilon_m} e^* \quad (16)$$

Equation (16) states that a strengthening of U.S. dollar with respect to Brazilian real decreases U.S. domestic production as American softwood plywood will be less competitive relative to Brazilian softwood plywood.

Finally, the reduced form equation for U.S. domestic consumption of softwood plywood is obtained by substituting Equation (14) into Equation (6):

$$Q_d^* = \frac{\eta_{us}\varepsilon_m}{\eta_m - \varepsilon_m} e^* \quad (17)$$

Equation (17) states that a strengthening of U.S. dollar with respect to Brazilian real increases U.S. consumption of softwood plywood.

2.2. Welfare formulas

With the fall in the price of Brazil softwood plywood due to the appreciation of U.S. dollar with respect to Brazilian real, the welfare effect of the U.S. dollar's strengthening can be approximated using the following formulas:

$$\Delta CS_{us} = -P_{us}^0\, Q_d^0\, P_{us}^*(1 + 0.5\, Q_d^*) \quad (18)$$

$$\Delta PS_{us} = P_{us}^0\, Q_s^0\, P_{us}^*(1 + 0.5\, Q_s^*) \quad (19)$$

$$\Delta PS_{br} = P_{us}^0\, P_{us}^*(Q_s^0\, Q_s^* - 0.5 Q_d^0 Q_d^*) \quad (20)$$

where ΔCS_{us}, ΔPS_{us}, and ΔPS_{br} are the change in the surpluses of U.S. consumers, U.S. producers, and Brazil producers.

3. Model calibration

In general, to estimate the reduced form elasticities in Equations (13)-(17), the parameters are taken from literature. Table 2.1 summarizes the reduced form elasticities expressions and corresponding equations.

Table 2.1. Summary of reduced form equations expressions

Endogenous variable	Reduced form elasticity Expression	Equation
P_{br}^*	$\frac{\eta_m}{\eta_m - \varepsilon_m} > 0$	(13)
P_{us}^*	$\frac{\varepsilon_m}{\eta_m - \varepsilon_m} < 0$	(14)
Q_m^*	$\frac{\varepsilon_m \eta_m}{\eta_m - \varepsilon_m} > 0$	(15)
Q_s^*	$\frac{\varepsilon_{us} \varepsilon_m}{\eta_m - \varepsilon_m} < 0$	(16)
Q_d^*	$\frac{\eta_{us} \varepsilon_m}{\eta_m - \varepsilon_m} > 0$	(17)

Haim et al. (2014) estimate the short-run own price elasticity of U.S. demand for softwood plywood for (1) new housing construction, (2) repair and remodeling, and (3) nonresidential construction at -0.29, -0.20, and -0.06, respectively. Haim et al. (2014) also estimate the long-run own price elasticity of demand for softwood plywood for the same construction sectors at -3.67, -0.88, and -0.26, respectively. Hence, we set the values of -0.18 and -1.60 for η_{us} for the short run and long-run, respectively. Assogba and Zhang (2021a) estimate the short-run and long-run U.S. supply elasticity for softwood plywood at 0.25 and 0.98, respectively. As this is the only paper on U.S. softwood plywood supply, we set the values of ε_{us} in the short-run and long-run at 0.25 and 0.98, respectively. The short-run and long-run values of U.S. import demand elasticity for Brazilian softwood plywood—η_m based on Assogba and Zhang (2021b) are set at -0.53 and -2.27, respectively. Using Equation (14), the value of the short-run export supply elasticity for softwood plywood from Brazil (ε_m) is set at 0.02.[5] Note

[5] From Equation (14), $\varepsilon_m = (PTE * \eta_m)/(1 + PTE)$. Yang (1997) estimates U.S. PTE for plywood at -0.03.

that the value of ε_m implies that U.S is a major importer of plywood from Brazil with a significant market power on Brazil's export market for plywood. We relax this assumption by using alternative values for ε_m (short-run: 0.44, long-run: 1.02 and 2.04) which implies a more elastic export supply function for Brazil's plywood. In 2008, for example U.S. total import of softwood plywood from Brazil accounted for only 22% of Brazil total exports of plywood (de Camargo, 2008). Table 2.2. presents the model parameters used for the estimation of the reduced form elasticities in the deterministic approach.

Table 2.2. Definition and baseline values for model parameters

Item	Definitions	Value	
		SR	LR
η_{us}	U.S. demand elasticity for softwood plywood	-0.18	-1.60
η_m	U.S. import demand elasticity for Brazilian softwood plywood	-0.53	-2.27
ε_{us}	U.S. domestic softwood plywood supply elasticity	0.25	0.98
ε_m	Brazil export supply elasticity for softwood plywood to U.S.	0.02 or 0.44[a]	1.02[a] or 2.04[a]
P_{us}^0	Initial equilibrium price of softwood plywood in U.S. ($\$/m^3$)	414.99[b]	
Q_s^0	Initial U.S. domestic production of softwood plywood (m^3)	8080935[c]	
Q_d^0	Initial U.S. consumption of softwood plywood (m3)	8126678[d]	

Note: SR and LR refer to Short-run (one quarter) and Long-run (four quarters).
[a] Under the assumption that Brazil export supply curve is relatively elastic.
[b] The average price of southern pine plywood and western fir plywood in 2010 is used as proxy. Source: Random Lengths 2015.
[c] U.S. domestic production of softwood plywood in 2010 is used as proxy. Source: Random Lengths 2015.
[d] U.S. consumption of softwood plywood in 2010 (computed as the sum of U.S. domestic production and U.S. imports of softwood plywood from Brazil) is used as proxy. U.S. import data is obtained from USDA Foreign Agricultural Service Global Agricultural Trade System (GATS) database (accessed on May 22, 2021).

The above approach assumes that the model parameters are known with certainty. However, to account for the uncertainty associated with the above parameters, the reduced form elasticities are also estimated using a stochastic simulation. Particularly, in the stochastic approach, the U.S. price elasticity of demand, supply, and Brazil export supply elasticities for softwood plywood are assumed to follow a triangular distribution with minimum, most likely, and maximum values given in braces{.}:

Short-run: ε_m = {0.02, 0.23, 0.44}, η_{us} = {-0.30, -0.18, -0.06}, ε_{us} = {0.20, 0.25, 0.75}, η_m = -0.53,

Long-run: ε_m = {1.02, 1.53, 2.04}, η_{us} = {-3.67, -1.60, -0.26}, ε_{us} = {0.80, 0.98, 1.50}, η_m = -2.27

The remaining parameters are assumed to be known with certainty.

4. Results and discussions

Table 2.3 presents the estimates of the reduced form elasticities using the deterministic approach of the exchange rate appreciation on the price of softwood plywood to U.S. consumers, Brazil producers, U.S. imports from Brazil, U.S. domestic production and consumption.

Table 2.3. Reduced form elasticities of exchange rate effect on U.S.-Brazil plywood market using the deterministic approach.

Endogenous variable	ε_m^{SR} = 0.02	ε_m^{SR} = 0.44	ε_m^{LR} = 1.02	ε_m^{LR} = 2.04
	SR	SR	LR	LR
P_{br}^*	0.96	0.54	0.69	0.52
P_{us}^*	-0.04	-0.46	-0.31	-0.48
Q_m^*	0.02	0.24	0.70	1.07
Q_s^*	-0.01	-0.11	-0.30	-0.46
Q_d^*	0.01	0.08	0.50	0.76

Notes: SR and LR refer to short-run and long-run.

The results in Table 2.3 show that the pass through to U.S. domestic consumers of softwood plywood associated with the dollar's appreciation (with respect to Brazilian real) is more perceptible, the more elastic is Brazil export supply curve and the longer is the time horizon. Particularly, with a relatively inelastic export supply curve (ε_m^{SR} = 0.02), an appreciation of U.S. dollar of 1% with respect to Brazilian real is associated with a decrease of U.S. consumers price of plywood by 0.04%. By contrast, with a more elastic Brazil export supply curve (ε_m^{SR} = 0.44), an appreciation of U.S. dollar of 1% with respect to Brazilian real is associated with a fall of U.S. consumers price of plywood by 0.46%. This is consistent with the literature on exchange rate pass through which suggests that the more elastic is the export supply curve, the larger is the pass-through elasticity to domestic consumers (Xie et al. 2008). Similarly, the longer is the time horizon, the larger is the pass-through elasticity. As for Brazilian producers, with a relatively inelastic export supply curve (ε_m^{SR} = 0.02), the appreciation of 1% of U.S. dollar with respect to Brazilian real increases the price of plywood received by Brazilian producers by 0.96%. By contrast when Brazil export supply curve becomes more elastic (ε_m^{SR} = 0.44), a 1% strengthening of U.S. dollar with Brazil real increases the price received by Brazilian producers only by 0.54%. This suggests that the absorption in the price of Brazilian plywood of U.S. dollar's strengthening with respect to Brazil real is smaller when U.S. faces a relatively elastic export supply curve.

The quantity effect of U.S. dollar's appreciation on U.S. imports and domestic production of softwood plywood is similar to the price effect. That is, the more elastic is Brazil export supply curve, the larger is U.S. imports of softwood plywood from Brazil. This is because, as the price of Brazilian plywood becomes cheaper due to the increase in the exchange rate pass through, U.S. demand for Brazilian plywood increases. As consequence, U.S. made plywood becomes less competitive which causes U.S. domestic production of plywood to fall. Particularly, when Brazil's export supply curve is relatively inelastic (ε_m^{SR} = 0.02), the appreciation of U.S. dollar of 1% with respect to Brazilian real is associated with an increase of U.S. imports by 0.02%. By contrast, when U.S. faces a relatively elastic export supply curve from Brazil (ε_m^{SR} = 0.44), an appreciation of U.S. dollar of 1% is associated with an increase of U.S. imports by 0.24%. Note that this increase in U.S. imports of plywood from Brazil is larger in the long-run than in the short-run. The negative effect on U.S. domestic production of U.S. dollar's strengthening with respect to Brazilian real increases with larger export supply curve from Brazil. Particularly, while the reduction of U.S. domestic production is negligeable with an inelastic export supply curve from Brazil (ε_m^{SR} = 0.02), the size of loss becomes important with a relatively elastic export supply curve from Brazil. This is because, when U.S. faces a relatively inelastic export supply curve from Brazil, the fall in price to U.S. consumers of Brazilian plywood associated with U.S. dollar's strengthening is smaller which induces a negligeable substitution effect between Brazilian and U.S. domestic plywood. U.S. consumption of softwood plywood increases, the larger is the pass-through effect of U.S. dollar's strengthening with respect to Brazilian real.

Table 2.4 presents the estimates of the reduced form elasticities using the stochastic approach. The mean and 95% confidence limits are generated from 1,000 random draws. The results in the stochastic approach are similar to those in the deterministic approach. That is the longer is the time horizon and/or the larger is the export supply elasticity of Brazil plywood, the larger is the price fall of Brazilian plywood to U.S. consumers when the U.S. dollar appreciates. Thus, U.S. import of softwood plywood from Brazil increases while U.S. domestic production of plywood falls. This implies that the loss in the competitiveness of U.S. domestic plywood increases following a strengthening of U.S. dollar with respect to Brazilian real when U.S. faces a relatively elastic export supply curve from Brazil.

Table 2.5 presents the welfare distribution following U.S. dollar's appreciation between 2010 and 2019 using the results from the stochastic simulation. The results show that although, the strengthening of U.S. dollar during the period 2010-19 had a detrimental effect on U.S. domestic producers of softwood plywood, the loss to U.S. producers was smaller than the gain to

U.S. consumers. Particularly both the short-run and long-run stochastic parameters show that the net welfare effect of U.S. dollar strengthening is positive. Using short-run parameters, the gains and the losses to U.S. consumers and producers are estimated per year at US $ 1.29 billion and US $ -1.15 billion, respectively. Similarly, the U.S. dollar's appreciation during the period 2010-19 with respect to Brazilian real was beneficial for Brazilian producers of softwood plywood both in the short-run and long-run. The increase in Brazilian producers' welfare using the short-run and long-run parameters are estimated at US $ 0.23 billion and US $ 1.66 billion, respectively.

Table 2.4. Reduced form elasticities using the stochastic simulation.

Endogenous variables	5% limit	Mean	95% limit
Short-run			
P_{br}^*	0.57	0.70	0.88
P_{us}^*	-0.43	-0.30	-0.11
Q_m^*	0.06	0.16	0.23
Q_s^*	-0.29	-0.13	-0.03
Q_d^*	0.03	0.05	0.06
Long-run			
P_{br}^*	0.54	0.60	0.67
P_{us}^*	-0.46	-0.40	-0.33
Q_m^*	0.75	0.91	1.05
Q_s^*	-0.65	-0.45	-0.29
Q_d^*	0.25	0.71	1.08

Table 2.5. Welfare effects based on stochastic simulations.

Period: 2010-19	Short run (US $ billion)	Long run (US $ billion)
Welfare distribution for U.S.		
ΔCS_{us}	1.29	2.41
ΔPS_{us}	-1.15	-1.20
Net Surplus	0.14	1.21
Welfare distribution for Brazil		
ΔPS_{br}	0.23	1.66

Conclusion

In this study, we investigate the effect of U.S. dollar's appreciation on U.S.- Brazil softwood plywood market and the welfare distribution of U.S. dollar strengthening with respect to Brazilian real during the period 2010-19. The Muth type modelling approach is adopted. The price and quantity effects of U.S. strengthening is estimated using both a deterministic and stochastic approaches. Particularly in the stochastic approach the parameters are assumed to follow a triangular distribution with 1,000 random draws. The results both in the short-run and long-run show that the pass-through effect of U.S. dollar's appreciation is more perceptible to U.S. consumers when U.S. faces a relatively elastic export supply curve from Brazil. Conversely, the absorption effect of U.S. dollar's strengthening is more perceptible to Brazilian producers when U.S. faces a relatively inelastic export supply curve from Brazil. Using the stochastic approach, a 1% appreciation of U.S. dollar with respect to Brazilian real is associated in the short-run with a fall in the price of Brazilian plywood to U.S. consumers by 0.30%. Similarly, a 1% strengthening of U.S. dollar with respect to Brazilian real is associated with an increase in the price of softwood plywood received by Brazilian producers by 0.70% in the short-run. In the long-run, a 1% strengthening of U.S. dollar with respect to Brazilian real is associated with a fall in the price of Brazilian plywood to U.S. consumers by 0.40%. Similarly, in the long-run, a strengthening of U.S. dollar of 1% is associated with an increase in the price received by Brazilian plywood producers by 0.60%. The quantity effects of U.S. dollar strengthening with respect to Brazilian real are similar to the price effects. Particularly, the appreciation of U.S. dollar with respect to Brazilian real is associated with an increase in U.S. imports of Brazilian plywood in the long-run. The fall in U.S. domestic production is also larger in the long-run. Overall, the U.S. dollar's strengthening with respect to Brazilian real during the period 2010-19 had a net positive welfare effect on U.S.-Brazil softwood plywood market. Although, the U.S. dollar’s strengthening with respect to Brazilian real had a detrimental effect on U.S. domestic producers of softwood plywood, the gains to U.S. consumers more than offset the losses to U.S. producers. As for Brazilian producers, they benefited from dollar’s appreciation with an increase in producers’ surplus. However, the conclusions of this study have to be interpreted with cautions. Particularly, we assumed that U.S. domestic softwood plywood and Brazilian plywood are perfect substitutes with a competitive market clearing. Further research could also investigate the effect of exchange rate movements on softwood plywood under commodities differentiation framework.

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Chapter 3: The Conservation Reserve Program and Timber Prices in the U.S. South

Abstract

In this study we estimate the effect of the Conservation Reserve Program (CRP) on the prices of standing timber in U.S. South using a reduced form model. The results of the estimation based on time series data covering the period between 1960 and 2019 suggest that CRP had a negative effect on the prices of standing timber in the long run. This statistically significant negative effect is estimated at -15.4 percent, which may partially explain why standing timber prices in the U.S. South have stayed low for more than a decade since the great financial crisis of 2008.

Key words: Conservation Reserve Program, timber price, supply, reduced form model, long-run.

1. Introduction

The Conservation Reserve Program (CRP) established by the Farm Bill of 1985 represents the largest environmental conservation program in the U.S. (Wallander et al. 2017). It provided annual payments to farmers who voluntary remove environmentally sensitive land from crop production to establish environmentally friendly cover (e.g., planting grasses or trees) for ten to fifteen years (Feng et al. 2005; Parks and Schorr 1997). Not only CRP aimed at land conservation, it was also used as a tool to control the oversupply of crop on the market following the booming of the U.S. agricultural sector in the 1970s and 1980s.

While the ecological and environmental benefits of CRP (e.g., Drum et al. 2015; Feather et al. 1999; Feather and Hellerstein 1997; Johnson and Igl 1995), farmers participation in the program (Jang and Du 2018; Wallender et al. 2017; Jacobs et al. 2014; Chang and Boisvert 2009; Feng et al. 2005; Plantinga et al. 2001), CRP's effect on slippage (Roberts and Bucholtz 2005; Wu 2005; Wu 2000), the economic incidence of CRP on farming industry (e.g., Brown et al. 2019; Wu and Lin 2010; Ryberg et al. 1991; Martin et al. 1988), the relationship between lands enrollment into CRP and the price of annual crops (e.g., Secchi and Babcock 2015; Hellerstein and Malcolm 2011), the post CRP land utilization (Onianwa et al. 1999; Barnes et al. 2020), and the cost effectiveness of the program (Miao et al. 2016; Baerenklau 2005; Kirwan et al. 2005; Khanna et al. 2003) have received much attention from scientists, we know of no published study in the agricultural and forest economics literature that investigated the effect of CRP on timber prices.

Standing timber is a capital asset, and the long-run supply of timber could be a couple of decades or more in North America (Zhang and Pearse 2011). In U.S. South, standing softwood timber often has a rotation age of 25 to 35 years with planted forests having a younger rotation age than natural forests. Thus, an increase in long-run timber supply induced by subsidized tree planting programs such as CRP in the later 1980s and earlier 1990s could only show its impact on timber prices a couple of decades latter. Yet, this linkage has not been established theoretically and demonstrated empirically in literature. Because the implementation of CRP in U.S. South mainly involved tree planting and especially increased the acreage of softwood forest acreages (Li and Zhang 2007; Butler and Wear 2013) and because U.S. South is the wood basket for the country, it provides a unique opportunity to investigate the linkage between long-run timber supply quantity and prices. Related, empirical evidence shows that landowners who planted trees often intent to keep tree cover and do not revert the land to crop production after CRP contracts are over (Onianwa et al. 1999).

In this study, we investigate the impact of CRP and other government cost-share programs such as Stewardship Incentives Programs and Forest Incentive Programs on long-run timber prices with the understanding that CRP was the most dominant government subsidy program related to tree planting in the South in the late 1980s and early 1990s. Therefore, CRP actually covers other programs as well here. The estimation of CRP effect on timber prices in U.S. South is of great importance for three reasons. First, we fill in a gap in the literature. Second and more importantly, our results show that the estimated incidence or impact of this program on landowners may be inflated because an increase in timber supply eventually leads to lower timber prices in the long-run. In other words, CRP may have unintended consequence in reducing timber prices a couple of decades later. As cost-share programs as a conservation tool become a common practice around the world, this calls for research on the economic incidence of these programs. Finally, U.S. South is the largest timber basket in the country and the world as it accounts for 60 percent of U.S. timber supply and some 16 percent global industrial roundwood and as some 86 percent of forest lands are family-owned. Thus, any meaningful change in timber prices would have a significant impact on the economic welfare of millions of family forest owners (Dezember 2018), although the decrease in the prices of timber may be a welfare increasing for forest products manufacturers and consumers of forest products. The results of this study could certainly help explain to landowners and policy makers why timber prices in the South have stayed low for more than a decade since the great financial crisis of 2008.

2. Theoretical model

The theoretical model developed in this study is similar to the model by Brannlund et al. (1985), Newman (1987), and Berck and Bentley (1997). Suppose that the price of sawtimber is determined by market forces, and that the intersection of demand and supply determines the equilibrium price of sawtimber on the market. Sawtimber market is assumed to, and proven to, be competitive (Murray 1995a, 1995b). Firms use sawtimber along with other inputs to produce goods (e.g., lumber) that maximize their profits. Let's define by π_{it}, the profit function of firm *i* in period *t*. The optimization problem of firm *i* can be expressed as follows:

$$\pi_{it}(p_{lmt}, p_{Lt}, p_{Kt}, p_{St}) = MAX_{W_{it}, K_{it}, S_{it}}\{ p_{lmt} \ f_{it}(L_{it}, K_{it}, S_{it}) - c_t' x_{it}\} \quad (1)$$

where f_{it} is the quantity of output produced by firm *i* in period *t* sold on the market at a unit price of p_{lmt}; L_{it}, K_{it} and S_{it} are respectively the quantity of labor, capital, and sawtimber used by firm *i* in period *t* for the output production; $f_{it}(L_{it}, K_{it}, S_{it})$ means that the quantity of output produced is a function of inputs used in the production; c_t' ($c_t' = [p_{Lt}, p_{Kt}, p_{St}]$) is a vector of the unit costs of inputs used by firm *i*; p_{Lt}, p_{Kt},, and p_{St}, are the unit price of labor, capital, and sawtimber respectively; and $x_{it}'= [L_{it}, K_{it}, S_{it}]$ is a vector of inputs quantity used by firm *i* in period *t*.

Because firm i is a price taker under the assumption that each firm is small enough to influence the market price of outputs, p_{lmt} is exogenously determined. Applying Hotelling's lemma, the derived demand curve of firm *i* for sawtimber is determined by differentiating its profit function with respect to sawtimber stumpage price:

$$\frac{\partial \pi_{it}}{\partial p_{St}} = Q_{di}\ (p_{lmt}, p_{Lt}, p_{Kt}\ , p_{St}) \qquad (2)$$

where Q_{di} denotes the derived demand for sawtimber by firm i, a function of output and input prices.

Assuming that U.S. South comprises *N* firms maintaining similar production functions and facing the same prices, the derived demand for sawtimber in the region can be obtained by aggregating individual firms' derived demands (Newman 1987). Thus,

$$Q_d\ (p_{lmt}, p_{Lt}, p_{Kt}\ , p_{St}) = \sum_{i=1}^{N} Q_{di}\ (p_{lmt}, p_{Lt}, p_{Kt}\ , p_{St}) \qquad (3)$$

Sawtimber stumpage supply in the South is the horizontal summation of all producers' supply curves. In the forest economics literature, it is derived either by considering the aggregation method which assumes identical production functions facing producers (Brannlund et al. 1985; Majumdar et al. 2010) or a simplified supply function in the case that timber producers face different productions functions (Newman 1987; Polyakov et al. 2005). In U.S. South, Newman (1987) noted the heterogeneity among timber producers (e.g., variability in management structure, government subsidies, production decisions, etc.) and recommended the use of a simplified supply function still accounting for the stumpage production costs and returns. This latter approach is adopted here.

Following Newman (1987) and Zhang and Pearse (2011), the sawtimber supply curve in the South is defined as a function of supply factors, namely, the sawtimber stumpage price, p_{St}, production costs which are represented approximately by standing timber inventory (Inv_t), and the policy of interest—the Conservation Reserve Program and other tree-planting subsidy programs (CRP_t):

$$Q_S = Q_S(p_{St}, Inv_t, CRP_t) \tag{4}$$

From equations (3) and (4), the market clearing conditions is specified:

$$Q_d\ (p_{lmt}, p_{Lt}, p_{Kt}\ , p_{St}) = Q_S(p_{St}, Inv_t, CRP_t) \tag{5}$$

Then solving equation (5) for the price of sawtimber on the market, the reduced form model for the commodity price can be expressed as:

$$p_{St} = f\ (p_{lmt}, p_{Lt}, p_{Kt}, Inv_t, CRP_t) \tag{6}$$

The reduced form model for the equilibrium price includes both demand and supply factors and has been used in the literature to model the prices of forest products (Brannlund et al. 1985; Newman 1987; Berck and Bentley 1997; Parajuli et al. 2016).

CRP_t and Inv_t are supply increasing factors, that is an increase in these factors is expected to induce an increase in sawtimber market supply. Hence, these factors are expected to have a negative effect on the price of sawtimber $\left(\frac{\partial p_{St}}{\partial CRP_t} < 0,\ \frac{\partial p_{St}}{\partial Inv_t} < 0\right)$. The increase in the price of the final product from sawtimber—lumber (p_{lmt}) is expected to have a positive impact on the supply of sawtimber. This increases the derived demand for sawtimber, and consequently increases sawtimber price $\left(\frac{\partial p_{St}}{\partial p_{lmt}} > 0\right)$. Note that the inventory variable is current and that the CRP variable must be lagged for 22-26 years to reflect the time for trees to grow into sawtimber in the South.

3. Empirical Model and Data

3.1. Empirical specification and variables

To estimate the effect of CRP on the price of sawtimber, the reduced form Equation (6) is assumed to be log linear in parameters. Hence, Equation (6) can be rewritten as:

$$\ln p_{St} = \beta_0 + \beta_1 \ln p_{lmt} + \beta_2\, lnp_{Lt} + \beta_3 \ln p_{Kt} + \beta_4 \ln Inv_t + \beta_5 \ln CRP_t + z'\delta + u_t \qquad (7)$$

where z is a vector of additional control variables and δ is a vector of parameters to be estimated, similar to $\beta_0 \ldots \beta_5$. The control variables include indicator variables for the great recession of 2008 (REC2008) and the Soil Bank Program (SBP_t), which took place between 1956 and 1962 and had the same impact as CRP on tree planting in the South.

The REC2008 variable is included because the great recession of 2008 had a negative effect on the U.S. economy and generated a significant decrease in the demand of timber and wood related products (Keegan et al. 2011, Ince and Nepal 2012). The objective of the Soil Bank Program was to reduce the surplus of farms commodities by incentivizing farmers to convert their farm lands into conservation land (Schmid 1958), similar to that of CRP. Again, the potential effect of Soil Bank Program on sawtimber stumpage prices is expected to be 22-26 years later when trees reach their merchantable ages. For this study, we use a time lag of 24 years after subsidized tree planting, although the results for other lags are similar to those reported below. Table 3.1 summarizes the variables definition.

Equation (7) is estimated using time series econometrics. The time series econometrics is used to avoid spurious regression as most of the variables are time series. Traditional regression techniques (e.g., ordinary least squares) require covariance stationary variables (Cryer and Chan 2008). A variable is covariance stationary if its first and second moments are finite and time invariant (Ogaki et al. 2005). Test of stationarity using the Augmented Dickey Fuller (ADF) test is implemented on each time series variables. The optimal number of lag variables included in the ADF test is selected using Schwarz's Bayesian information criteria (SBIC). Furthermore, the cointegration relationship of the time series variables is explored.

3.2. Data

The data used for the empirical model covers the period between 1960 and 2019. Table 3.1 presents the variable definition. The price of pine sawtimber in the U.S. South is used as a measure of timber prices in the region. The price variable has been adjusted to the common base year of 2019 by means of the Producer Price Index for all commodities (PPIACO, base 2019 = 100) obtained from the website of the St. Louis Federal Reserve Bank (FRED 2020). The sawtimber price data is obtained from Timber-Mart South, Inc. (various years) for 1977 and 2019 and Ulrich (1990) who used sawtimber prices from Louisiana before 1977. Figure 3.1 presents the real price of sawtimber dynamic in U.S. South from 1960 to 2019.

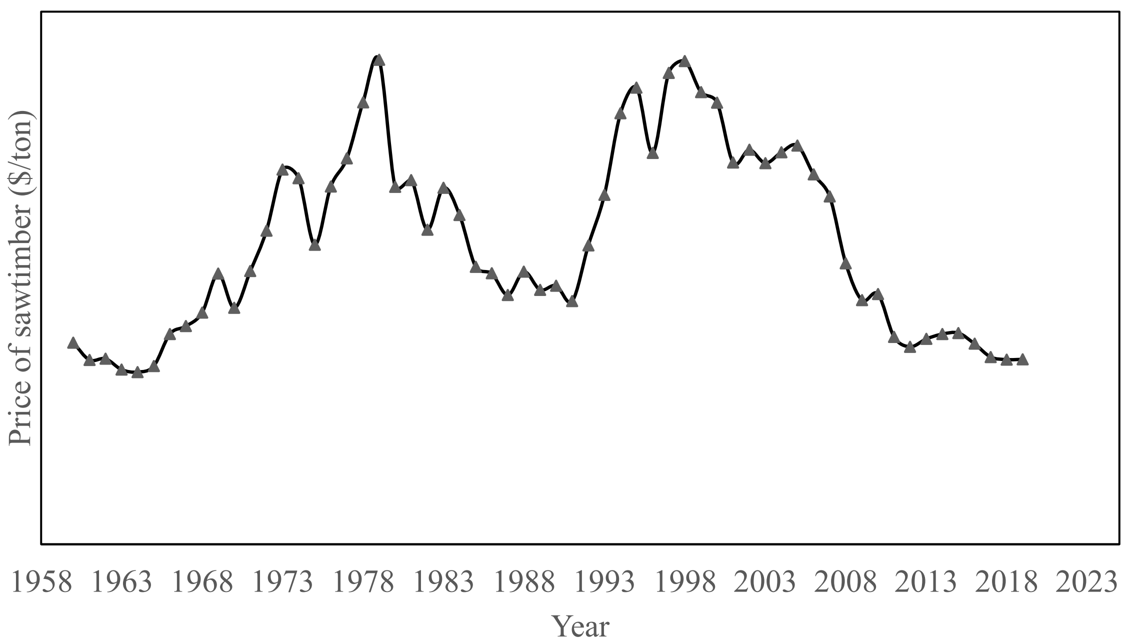

Figure 3.1: Real price of sawtimber for the period 1960 to 2019

Table 3.1. Variable definition

Variable name	Description	Sources
p_{St}	Real Price of Sawtimber stumpage (dollar per ton) from 1960 to 2019	Timber-Mart South, Inc. (various years) for 1977 and 2019; Ulrich (1990) before 1977.
p_{Lt}	Cost of labor — average hourly earnings in the U.S. wood products industry (CEU313210032), adjusted by U.S. Producer Price Index for all commodities (PPIACO) (2019 = 100)	FRED (2020)
p_{Kt}	Cost of Capital — U.S. bank prime loan rate (DPRIME) (2019 =100)	Same as above
p_{lmt}	Unit price of lumber — U.S. producer price index for lumber (WPU081) (2019 = 100)	Same as above
Inv_t	Inventory — Net volume of sawtimber on timberland in USA (in thousand cubic meter)	Oswalt et al. (2019).
CRP_t	A dummy variable for the effect of CRP (taking the value of 1 between 2010 and 2019)	Constructed by the authors
SBP_t	A dummy variable for the Soil Bank Program (taking the value of 1 between 1980 and 1986)	Constructed by the authors
$subsidized\ tree\ acres_t$	Subsidized tree planting acres in U.S. south	Zhang (2004); Li and Zhang (2007)
$non\ subsidized\ tree\ acres_t$	Non-subsidized tree planting acreages in U.S. south	Same as above
$Rec2008_t$	A dummy variable for the great recession crisis of 2008 (taking the value of 1 between 2008 and 2009)	Constructed by the authors

* subscripts denote year t

The U.S. producer price index for lumber (WPU081, base 2019 = 100) is used as measurement for the price of the final goods from sawtimber. The data is also obtained from FRED (2020). The average hourly earnings in the U.S. wood products industry (CEU313210032) is used as approximation for the cost of labor and adjusted for inflation by means of the Producer Price

Index for all commodities (FRED 2020). Note that the wage variable was only available for the period 1990-2019. The data for the missing period is computed by using the predicted values of the regression of the average hourly earnings in the wood products on the average hourly earnings in the U.S. manufacturing industry. The data on the average hourly earnings in the U.S. manufacturing industry (CES300000008) and U.S. wood products industry are obtained from FRED (2020) and Bureau of Labor Statistics (BLS 2020), respectively.

The bank prime loan rate (DPRIME, base 2019 =100) is used as approximation for the cost of capital in the timber-based industries. The bank prime loan rate has been used in the literature as measurement of cost of capital in the timber-based industry (Haslett 1986; Majumdar et al. 2010). The data for this variable is also obtained from FRED (2020).

The data on the timber inventory in U.S. South is obtained from Oswalt et al. (2019). Because the inventory data are reported for specific years with an approximate 5-year interval, the data for the years not reported are estimated following Newman (1987), which is current timber inventory is equal to the timber inventory in previous year, plus estimated growth, and minus harvesting and loss due to natural disasters. The data on subsidized and non-subsidized tree planting acreages are from Zhang (2004) and Li and Zhang (2007). As Figure 3.2 shows, there were three periods (1956 to 1963, 1986 to 1993, and 1998) when tree planting was heavily subsidized by the Soil Bank Program and CRP (in addition to other government subsidy programs), respectively, and the first two periods are relevant to this study. Subsequently, we use two approaches to account for the impact of Soil Bank Program and CRP. The first is to use two dummy variables to represent the 24-year lagged subsidized tree planting activities represented by these two programs. The other is to use lagged tree planting acreage variables by 24 years.

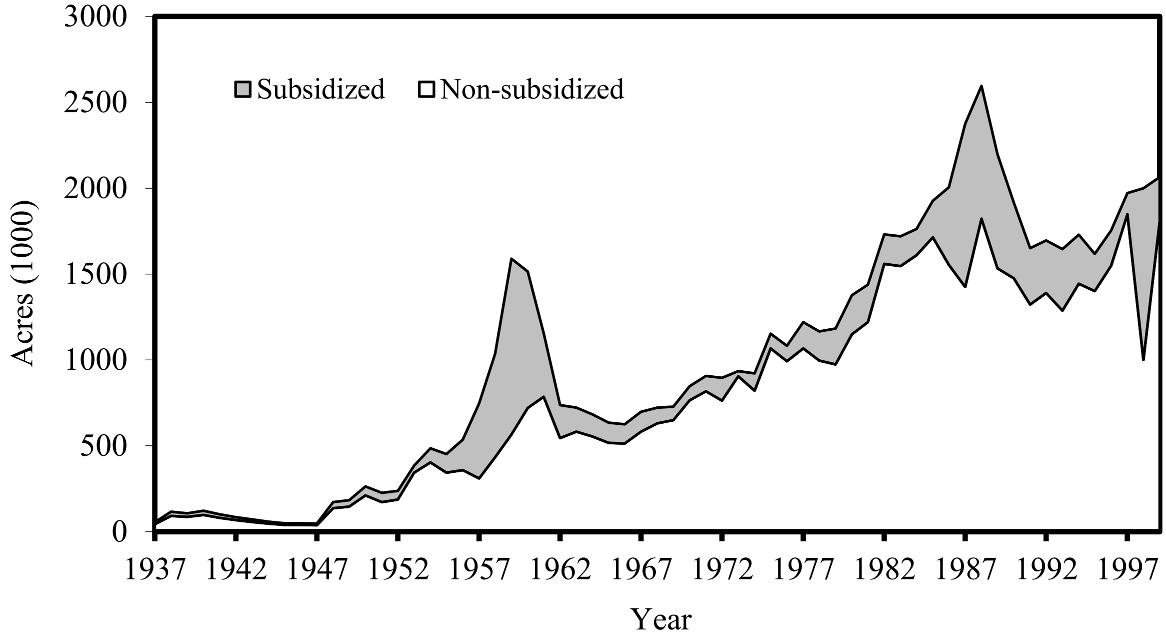

Figure 3.2: Subsidized and Non-subsidized Private Tree Planting in the U.S. South, 1937-1999

4. Results and Discussion

4.1. Unit Root Tests

The stationarity of each time series variable is examined using the Augmented Dickey Fuller (ADF) unit root test. The unit root test is conducted for equations with intercept only, and intercept and linear trend. Table 3.2 presents the optimal lags number and the unit root test on the time series variables. The optimal number of lags is selected using Schwarz's Bayesian information criteria (SBIC). The results suggest that all the five (5) time series variables are level non-stationary and first difference stationary.

Table 3.2. Summary of the unit root pretest on time series variables

Variables	Specification	Lags	ADFc	ADFc,t
p_{St}	log	0	-1.30	-1.13
	first difference in log	0	-7.09***	-7.43***
p_{Lt}	log	0	-1.71	-1.57
	first difference in log	0	-12.53***	-12.72***
p_{Kt}	log	2	-1.67	-2.21
	first difference in log	1	-7.01***	-7.04***
p_{lmt}	log	0	-1.77	-1.04
	first difference in log	0	-6.40***	-6.66***
Inv_t	log	0	-0.64	-2.63
	first difference in log	1	-5.76***	-5.69***

Notes: ADF_c and $ADF_{c,t}$ refer to ADF-t statistics when an intercept is included and when an intercept and time trend are included. *, **, *** The null hypothesis of unit root is rejected at the 10%, 5% and 1% level, respectively.

4.2. Cointegration results

The results in Table 3.2 show that all the time series variables are integrated of order 1. Hence, their cointegration relationship can be explored using Engle and Granger (EG) test of cointegration (Engle and Granger 1987). Time series variables integrated of order 1 are said to be cointegrated if their linear combination is stationary (Engle and Granger 1987). EG test of cointegration is a robust alternative to the multivariate Johansen tests for cointegration (MacKinnon 2010; Assogba and Zhang 2021). The EG test for cointegration is preferred over the multivariate Johansen test for cointegration because it makes fewer distributional assumptions and recommended for single equation reduced form model.

The serial correlation of the error terms is tested using Breusch-Godfrey statistics. The cointegration tests are conducted for equations under three assumptions (i) intercept only, (ii) intercept and linear trend and (iii) intercept and linear and quadratic trend (MacKinnon 2010). Table 3.3 presents the summary of the cointegration test regression at 5% critical values provided by MacKinnon (1990, 2010).

Table 3.3. Summary of Engle and Granger (EG) cointegration test regression

Assumptions	EG test statistics	Cointegration? (Yes/No)
Intercept only included in EG first step regression	-2.02	No
Allowing for deterministic trend in EG first step regression		
Intercept + linear trend	-2.03	No
Intercept + linear trend + quadratic trend	-2.06	No

*, **, *** The null hypothesis of no cointegration is rejected at the 10%, 5% and 1% level, respectively.

The results in Table 3.3 suggest no evidence of cointegration among the time series variables with the price of sawtimber being the response variable. Hence, an error correction term can be ignored in the estimation of the reduced form equation with appropriate transformation of the time series variables (Granger 1981; Greene 2002; Mackinnon 2010).

4.3. Estimation of the Effect of CRP

Table 3.4 presents the estimation results of the reduced model. All the time series variables in the model are first differenced as the ADF test results show no evidence of level stationarity in the variables. The lag variables are included to control for the serial correlation in the error term. The goodness of fit of the model, as measured by the R^2 and the F-statistics, is relatively high. Further, the Bartlett's white noise test on the residuals of the model fails to reject the null hypothesis of white noise of the residuals. This suggests a satisfactory specification of the regression model. All the coefficients of the demand and supply factors in the regression model have the expected signs. The supply increasing factors (e.g., CRP, Soil Bank, and lagged inventory) have the expected negative sign on the sawtimber equilibrium prices. Similarly, the demand decreasing variable— Rec2008—has the expected negative sign on the sawtimber equilibrium prices. The positive sign of the variables labor and capital cost may suggest that sawtimber and these inputs are gross substitute inputs in the production of forest products from sawtimber stumpage. Finally, the results from the reduced form model suggest that CRP had a negative effect on the prices of sawtimber. This effect is estimated at -34.5 percent.[6]

[6] The percentage effect of CRP as measured by the dummy variable on the price of sawtimber is computed as the average of the ratio $(\hat{y}_1 - \hat{y}_0)/\hat{y}_0$ for the period 2010 to 2019. $\hat{y}_1$ and $\hat{y}_0$ are the predicted price of sawtimber with CRP and without CRP, respectively.

Table 3.4. Estimation results of the effect of CRP as measured by dummy variable.

Variables	Dependent variable = D $\ln p_{St}$	
	Coefficients	Standard Error
D $\ln p_{Lt}$	0.959[c]	(0.269)
D $\ln p_{Kt}$	0.196[c]	(0.067)
D $\ln p_{lmt}$	0.858[c]	(0.120)
D $\ln Inv_t$	0.407	(0.431)
D $\ln Inv_{t-1}$	-0.093	(0.440)
D $\ln p_{S,t-1}$	-0.100	(0.099)
CRP_t	-0.085[b]	(0.028)
SBP_t	-0.072[a]	(0.033)
Rec2008	-0.069	(0.063)
Constant	0.022	(0.017)
R-squared	0.65	
Adj R-squared	0.59	
Residuals' Bartlett's white noise test statistics	1.02	

Note: a, b, and c indicate significance at the 10%, 5% and 1% levels, respectively. The symbol "*D*" stands for the first difference of the variables. Standard errors are in parentheses.

Table 3.5 presents the estimation results of the effect of CRP as measured by the total tree acreages of government subsidies with a 24-year lag on the prices of sawtimber stumpage. The goodness of fit of the model, as measured by the R^2 and F-statistics, is relatively high as well. Further, the Bartlett's white noise test on the residuals of the model fails to reject the null hypothesis of white noise of the residuals—suggesting a satisfactory specification of the regression model. All variables have similar and consistent results as reported in Table 3.4. The result for the lagged acreage variable show that the CRP-induced tree planting had a negative effect on the price of sawtimber. By reducing the subsidized tree planting acreage induced by CRP to zero, we find that this effect is -15.4 percent.[7]

We believe that this result is reasonable given the facts that subsidized tree planting acreage accounted for 22.4 percent of all tree planting acreage in the U.S. South between 1986 and 1996 and some 40% of the timber inventory could be classified as sawtimber at age 24 (Fu 2014), that the inventory supply elasticity is empirically estimated to be around 0.8-1.2 (Newman 1987, Binkley 1988, Zhang and Pearse 2011), and that the demand elasticity for sawtimber stumpage in the U.S. South is estimated to be around -0.6 (Newman 1987). We also believe the estimated results in Table 3.5 is more accurate than those reported in Table 3.4 which uses the dummy variable approach, because the dummy variable approach catches all of the effects of tree plantings and treat tree planting equal in all the intervening years.

4.4. A Comparison

For comparison purpose, we estimated the reduced form Equation (7) using the specification and OLS approach by Newman (1987). Table 3.6 and 3.7 present the estimation results of the effect of CRP as measured by dummy variable and subsidized tree acreages, respectively. Model 1 shows the effect of CRP on the price of sawtimber when ignoring for the lag effects of both the inventory and the price of sawtimber. Model 2 shows the effect of CRP on the price of sawtimber when ignoring only for the effect of the lag variable of the price of

[7] The percentage effect of CRP as measured by subsidized tree acreages is computed as the average of the ratio $(\hat{y}_1 - \hat{y}_0)/\hat{y}_0$ for the period 2010 to 2019. $\hat{y}_1$ and $\hat{y}_0$ are the predicted price of sawtimber with and without tree acreages enrolled in CRP, respectively.

sawtimber. Model 3 differs from Model 1 and Model 2 by accounting for the lag effects of the variables inventory and the price of sawtimber, respectively.

Table 3.5. Estimation results of the effect of CRP as measured by subsidized tree acreages.

Variables	Dependent variables = D $\ln p_{St}$	
	Coefficients	Standard Error
D $\ln p_{Lt}$	0.750[c]	(0.270)
D $\ln p_{Kt}$	0.152[b]	(0.066)
D $\ln p_{lmt}$	0.846[c]	(0.121)
D $\ln Inv_t$	0.487	(0.455)
D $\ln Inv_{t-1}$	-0.262	(0.448)
D $\ln p_{S,t-1}$	-0.090	(0.099)
ln $subsidized\ tree\ acres_{t-24}$*	-0.042[b]	(0.016)
ln $non\ subsidized\ tree\ acres_{t-24}$ *	0.021	(0.018)
Rec2008	-0.060	(0.065)
Constant	0.061	(0.067)
R- squared	0.65	
Adj R-squared	0.58	
Residuals' Bartlett's white noise test statistics	0.90	

Note: a, b, and c indicate significance at the 10%, 5%, and 1% level, respectively. The symbol "*D*" stands for the first difference of the variables. Standard errors are in parentheses.

* The variables subsidized tree acres and non-subsidized tree acres are both level stationary as the ADF test rejects the null hypothesis of unit root at 1%.

Table 3.6. Estimation results of the effect of CRP as measured by the dummy variable using the linear model in Newman (1987)

Variables	Dependent variable $= \ln p_{St}$					
	Model 1		Model 2		Model 3	
	Coefficients	S.E.	Coefficients	S.E	Coefficients	S.E.
$\ln p_{Lt}$	0.927^{c}	(0.220)	0.923^{c}	(0.224)	0.660^{c}	(0.133)
$\ln p_{Kt}$	0.126	(0.075)	0.119	(0.078)	-0.005	(0.046)
$\ln p_{lmt}$	1.119^{c}	(0.174)	1.144^{c}	(0.186)	0.667^{c}	(0.118)
$\ln Inv_{t}$	-0.146	(0.489)	-0.067	(0.860)	0.152	(0.498)
$\ln Inv_{t-1}$			-0.365	(0.929)	-0.265	(0.538)
$\ln p_{S,t-1}$					-0.698^{c}	(0.070)
CRP_{t}	-0.319^{b}	(0.156)	-0.294^{b}	(0.131)	-0.043	(0.080)
SBP_{t}	0.011	(0.082)	-0.006	(0.081)	-0.014	(0.047)
Rec2008	0.092	(0.144)	0.016	(0.149)	0.017	(0.087)
Constant	-2.040	(6.105)	0.280	(7.022)	-1.952	(4.070)
R-squared	0.82		0.81		0.94	
Adj. R-squared	0.79		0.78		0.93	

Note: a, b, c indicates 10%, 5% and 1% level, respectively. Standard errors in parentheses.

Table 3.7. Estimation results of the effect of CRP as measured by subsidized tree acreages using the linear model in Newman (1987)

Variables	Dependent variable $= \ln p_{St}$					
	Model 1		Model 2		Model 3	
	Coefficients	S.E.	Coefficients	S.E.	Coefficients	S.E.
$\ln p_{Lt}$	1.128[c]	(0.307)	1.094[c]	(0.308)	0.651[c]	(0.174)
$\ln p_{Kt}$	0.172[b]	(0.078)	0.156[a]	(0.079)	-0.021	(0.047)
$\ln p_{lmt}$	1.376[c]	(0.194)	1.395[c]	(0.195)	0.750[c]	(0.122)
$\ln Inv_t$	-0.934[a]	(0.476)	-0.120	(0.922)	0.017	(0.506)
$\ln Inv_{t-1}$			-0.966	(0.938)	-0.304	(0.518)
$\ln p_{S,t-1}$					-0.717[c]	(0.066)
ln subsidized tree acres_{t-24}	0.014	(0.039)	0.007[b]	(0.039)	0.012	(0.022)
ln nonsubsidized tree acres_{t-24}	0.005	(0.090)	-0.006	(0.090)	-0.049	(0.050)
Rec2008	0.346[b]	(0.132)	0.331[b]	(0.133)	0.068	(0.077)
Constant	5.478	(6.880)	7.410	(7.128)	0.126	(3.968)
R-squared	0.79		0.79		0.94	
Adj. R-squared	0.76		0.76		0.93	

Note: a, b, c indicates 10%, 5% and 1% level, respectively. Standard errors in parentheses.

The results in the three models in Table 3.6 are consistent with high goodness of fits as measured by the R^2 and F-statistics compare to our model in Table 3.4. The signs of the coefficients of the demand and supply factors in all the three models are consistent with the results in our models in Table 3.4 with the exception of REC2008 and SBP variables. Further, the magnitudes of the effect of CRP on the price of sawtimber in Model 1 and Model 2 are similar to our estimate.

Particularly, the estimation results in Table 3.6 suggest that CRP decreased the price of sawtimber by -31.7 percent and -30 percent in models 1 and 2, respectively.

However, when accounting for both the lag effects of the variables inventory and price of sawtimber in Model 3 in Table 3.6, no evidence on the effect of CRP on the price of sawtimber is observed. Similarly, the estimation results of the effect of CRP as measured by subsidized tree acreages in Table 3.7 suggest no evidence of the effect the effect of CRP on price of sawtimber. These observed differences between the OLS approach by Newman (1987) and our approach could be explained by the ignorance of the order of integration of the time series variables in the regression model in Newman (1987). We believe that our approach is more reliable as we accounted for the order of integration of the time series variables in the regression model as opposed to Newman (1987) approach.

Conclusions

In this study we investigate the effect of the Conservation Reservation Program (CRP) on the equilibrium prices of sawtimber prices in the U.S. South. For this purpose, a reduced form model for sawtimber price is developed and estimated using data from 1960 to 2019. The time series econometrics that includes the tests of stationarity, the cointegration analysis and the linear model with stationary time series variables was used. The results of the estimated model suggest that CRP has a negative effect on the equilibrium prices of standing timber. This effect is estimated at -15.4 percent, all else being equal.

This result suggests that CRP induced a significant increase in the supply of sawtimber that depressed its equilibrium price after 24 years. Thus, the link between quantity supplied and prices in the long-term timber supply curve is evidently established. This result may partially explain why sawtimber stumpage prices have stayed low for so long after the 2008 great financial crisis—an oversupply of sawtimber in addition to low housing starts and other demand factors. This oversupply of sawtimber is an unintended consequence of CRP and other government subsidy programs implemented. While CRP was a direct welfare transfer to landowners who produce timber, conducting the economic analysis of CRP assuming a steady price would inflate the impact of such a transfer. On the other hand, CRP benefitted the forest product industry that uses sawtimber as input in the production of forest products and final consumers. A thorough economic analysis on government cost-share programs such as CRP, which have become popular around the world in order to curb environmental degradation, require consideration of all direct and indirect impacts.

Our results also have practical implications on landowners' decision to harvest timber. Because standing timber is a factory (a capital asset) as well as a product, many landowners and investors have used “store-value-on-the-stump” approach when timber market is down. This approach works best when timber market behaves like a random walk or any low prices exist only in the short-run, say 1-5 years. When timber market does not behave like this and stumpage prices do not recover or does not recover soon enough as in the last decade or more, the appropriateness of this approach needs to be scrutinized. Implicit in this approach to forego income today is the expectation that one can more than recover the forgone income in the future using the landowner’s cost of capital. In theory, the increase in value comes from a combination of biological growth and real price appreciation. Even in situations where prices in real terms recover and the accumulated volume can be harvested in the distant future, it is entirely possible that higher discount rates now or in the future could more than offset high level of projected income in the future.

Finally, the long-term nature of timber production means that timber market is inherently more difficult to forecast than annual or biannual crop products. With increasing number of forest landowners and institutional investments in timberland in the U.S. and elsewhere, a reliable long-term timber market forecast is more important than ever. Landowners and investors should help support future research on long-term trends of timber markets in various parts of the country and the world.

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