
The Effects of Exchange Rate Volatility on U.S. Forest Commodities Exports

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ABSTRACT. This article addresses the impact of exchange rate volatility on U.S. exports of four forest commodities. Exchange rate volatility is measured by the standard deviation of the growth rate of real effective exchange rate of the U.S. dollar. The nonstationarity of individual time series is explicitly taken into account by employing multivariate cointegration analysis and error correction models. The results show that exchange rate volatility has a negative impact on U.S. exports in the long term, but short-term dynamics vary for different commodities. A stable currency policy in the long run helps promote U.S. exports of forest commodities, although some commodities may benefit from exchange rate volatility in the short term. *For. Sci.* 49(5):807–814.

Key Words: International trade, chips, logs, pulp, time-series analysis.

FOREIGN EXCHANGE RATES have been highly volatile since the ending of fixed exchange rates in March 1973. According to the real effective exchange rate of the U.S. dollar constructed by the Federal Reserve Board of Chicago, the U.S. dollar appreciated by about 60% with respect to major currencies between 1978 and 1985; it then depreciated until 1988 when its value was close to its 1978 level. In the 1990s, the U.S. dollar was weak in the first half but became stronger in the second half. The elevated volatility and uncertainty of exchange rate movements have led researchers to investigate the impact of such movements on international trade.

A large number of studies have focused on the effect of exchange rate volatility on international trade flows since the adoption of a flexible exchange rate system. The standard hypothesis is that unexpected change in exchange rates affects the decisions made by risk-averse traders and thereby reduces trade. However, both theoretical and empirical contributions to the literature [1] fail to conclusively validate the hypothesis (Cote 1994, McKenzie 1999). For example, Chowdhury (1993) examines the impact of exchange rate volatility on the trade flows of the G-7 countries in the context of a multivariate error-correction model. The results indicate that the exchange rate

volatility has a significant negative impact on the volume of exports in each of the G-7 countries. Applying the similar techniques, Arize (1997a, 1997b), Arize et al. (2000), and Hassan and Tufte (1998) find similar results. On the other hand, Asseery and Peel (1991) provide empirical evidence that contradicts these findings.

These indeterminate results may be partially related to the fact that previous studies focus on aggregate trade flows, which can obscure commodity or sector-specific exchange rate effects (Anderson and Garcia 1989). With aggregate trade flows, it is implicitly assumed that the impact of exchange rate volatility is uniform among countries and for all commodities in terms of both direction and magnitude (McKenzie 1999). This study is motivated in part by the need for empirical evidence at the disaggregate level. Focusing on the exchange rate effect on specific commodities (i.e., forest commodities in this study) will not only enhance our understanding of the dynamics of the export market of forest commodities but also contribute to the literature of exchange rate volatility.

The objective here is to evaluate the effect of exchange rate uncertainty on U.S. exports of forest commodities. Exchange rates have been incorporated in trade studies of various forest commodities (e.g., Adams et al. 1986,

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Uusivuori and Buongiorno 1991, Alavalapati et al. 1997). No study, however, has explicitly examined the role of exchange rate volatility on forest commodities trade with the exception of Hannien (1999), who focuses on Finnish sawnwood exports to the United Kingdom. In this study, we examine U.S. exports of four forest commodities—wood chips (CHIP), softwood logs (LOG), dissolving wood pulp (DWP), and bleached sulphate wood pulp (BSWP) (as defined below). Cointegration analysis and an error-correction model are employed to build the long-term relationship and short-term dynamics between exports and their various determinants. The use of disaggregated trade data in evaluating the impact of exchange rate volatility on trade flows may be potentially beneficial for corporate trade strategy and policy formulation.

The remainder of this article is organized as follows. The next section discusses econometric methodology, followed by model specification and data. The final sections present the empirical results and draw some conclusions.

Methodology: Time Series Analysis

Not so long ago, time series data were treated similarly to cross-sectional data. Regressions on some nonstationary time series data can lead to spurious values of R -squared, DW , and t statistics, causing economists to conclude erroneously that meaningful relationships exist. As inference with OLS (Ordinary Least Square) is invalid for nonstationary data (Kennedy 1998, p. 263), the recognition that some economic variables, such as exports and their determinants, might be nonstationary integrated variables has led to research that has markedly changed and improved the way time series data is analyzed.

In this study, the Johansen cointegration analysis and vector error correction model (ECM) are employed. Our method starts with the vector autoregressive model (VAR). Consider a vector X_t with N nonstationary variables of interest [2], defined by a general polynomial distributed lag process as

$$X_t = \pi_1 X_{t-1} + \dots + \pi_k X_{t-k} + \varepsilon_t \quad (1)$$

where $t = 1, \dots, k$, and ε_t is an independently and identically distributed N dimensional vector with zero mean and variance-covariance matrix. It can be reformulated as

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \varepsilon_t \quad (2)$$

where

$$\Gamma_i = - \left(I - \sum_{i=1}^{k-1} \pi_i \right), \quad \Pi = - \left(I - \sum_{i=1}^k \pi_i \right)$$

and I is a $N \times N$ identity matrix.

Equation (2) contains information on both the short-run and long-run adjustments to changes in X_t via the estimates of $\hat{\Gamma}_i$ and $\hat{\Pi}$, respectively. The number of distinct cointegrating vectors (r) that exist among the variables of X_t is given by the

rank of Π . Π is an $N \times N$ matrix, and also defined as two of $N \times r$ matrices, α and β , such that

$$\Pi = \alpha \beta' \quad (3)$$

where α represents the speed of adjustment to disequilibrium, and β is a matrix of long-run coefficients. The existing of cointegrating relationships indicates that, although X_t is nonstationary, the linear combinations of $\beta' X_t$ are indeed stationary, and hence the columns of β form r distinct cointegrating vectors.

In general, the analysis of time series variables in the multivariate context involves three steps (Enders 1995, p. 266). The first step is to determine the integration order of individual time series, which is the prerequisite for cointegration analysis. The Augmented Dickey-Fuller (ADF) unit root test is often applied to test whether a time series has a unit root.

If the unit root tests reveal that the variables are integrated of the same order of one, the next step is to estimate a long-run equilibrium relationship using cointegration analysis. The typical method is the Johansen cointegration test (Johansen 1988, 1991, Johansen and Juselius 1990). For the Johansen method, the trace test and the maximum eigenvalue test are employed to determine the number of cointegrating vectors.

If the variables included are cointegrated, the final step is to investigate the dynamic behavior of the model by specifying and estimating a vector error correction model, which includes an error correction term. The correspondence between cointegration and the error correction model is formalized in the Granger Representation Theorem (Engle and Granger 1987). The size of the error correction term indicates the speed of adjustment of any disequilibrium towards a long-term equilibrium state.

Model Specification and Data

The effect of exchange rate volatility on U.S. forest commodities exports can be analyzed using an export demand model. For individual forest commodities, an export demand function is specified and estimated, with exchange rate volatility being one of the factors that affects export levels. Drawing upon the empirical literature in this area (Arize 1997a, 1997b, Chowdhury 1993), a long-run specification of the export demand equation in the flexible exchange rate environment can be expressed for each commodity as

$$Q_t = \beta_0 + \beta_1 P_t + \beta_2 G_t + \beta_3 V_t + e_t \quad (4)$$

where β s are coefficients, t indexes time, and all variables are expressed in natural logarithm.

Q_t is total U.S. export quantities of a specific forest commodity at time t . P_t is the ratio of U.S. export price to the world export price, which indicates the price competitiveness of that particular forest commodity from the United States in the world market. The world export price is the alternative price faced by prospective importers of the forest commodity from the United States and is measured as a weighted average of export prices of other major exporting countries in the world market. The variable G_t captures world demand condi-

tion and world effective purchasing power, and is measured by the trade-value-weighted average of the real gross domestic product (GDP) of major importers.

V_t is the exchange rate volatility that has been introduced indirectly by expressing all monetary variables in common dollar units. It is measured by a moving sample standard deviation of the growth rate of the real exchange rate as

$$V_t = \left[\frac{1}{m} \sum_{i=1}^m (\ln R_{t+i-1} - \ln R_{t+i-2})^2 \right]^{1/2} \quad (5)$$

where R is the real effective exchange rate of the U.S. dollar, and $m = 6$ is the order of the moving average. Work by Baba et al. (1992) provides the basis of employing this measure. Furthermore, this measure is similar to those employed in the exchange-rate volatility and international trade literature (e.g., Koray and Lanstapes 1989, Chowdhury 1993, Arize et al. 2000).

As far as the sign of each variable is concerned, it is expected that the price variable will be negative. The income variable could carry either negative or positive coefficients, depending on the relative sensitivity of the demand for traded goods and the production of the trade-substitute goods. In other words, rising global income often causes more demand for traded good, making the income variable positive. Rising demand could give rise to domestic substitutes and reduce the demand for traded commodities in some cases. For exchange rate volatility, if it deters the trade as hypothesized, it will carry a negative coefficient.

Four forest commodities are selected for the analysis. They are wood chips and particle (*CHIP*), softwood sawlogs, and veneer logs (*LOG*), dissolving wood pulp (*DWP*), and bleached sulphate wood pulp (*BSWP*). The United States has been among the largest exporters in the international market for these forest commodities, as re-

vealed by the export shares in Figure 1. Three quarterly series (Q_t , P_t , and G_t) for each of the four forest commodities selected plus the volatility measure V_t are constructed from the raw data. The sample covered 24 yr from 1978 to 2001 except *LOG* (1978–1988). Details about the data sources and construction are presented in Appendix I.

Empirical Results

The stationarity of each time series is first examined by the ADF test. The null hypothesis of unit root is rejected for all the series in the first difference but failed in the level as shown in Table 1. Therefore, all variables are integrated of order one. Next, the cointegration test and an error correction model are examined for each of the four forest commodities.

Cointegration Test and Long-Term Equilibrium

Since all the individual series (Q , P , G , and V) for the forest commodities are integrated of order one, it is possible to examine the existence of a long-term equilibrium between them using cointegration analysis. Johansen (1988, 1991) and Johansen and Juselius (1990) suggest starting the analysis by choosing an appropriate lag length for the traditional VAR model and then using the same lag length in cointegration analysis. The strategy adopted in specifying the lag length is based on calculating the log likelihood of the whole system and applying the Sims (1980) likelihood ratio (LR) test corrected for the degrees of freedom. The Akaike Information Criteria (AIC) criteria and Schwarz Criteria also can be used to evaluate the models. Based on the LR test, the lag length is specified as three for *CHIP*, one for *LOG*, four for *DWP* and *BSWP* in this study. For *CHIP* and *DWP*, a trend is allowed in the cointegration regression since the null hypothesis of an intercept in the cointegrating vectors against the alternative of a linear trend in the variables is rejected. For *LOG* and *BSWP*, only constants are included in the cointegration regression.

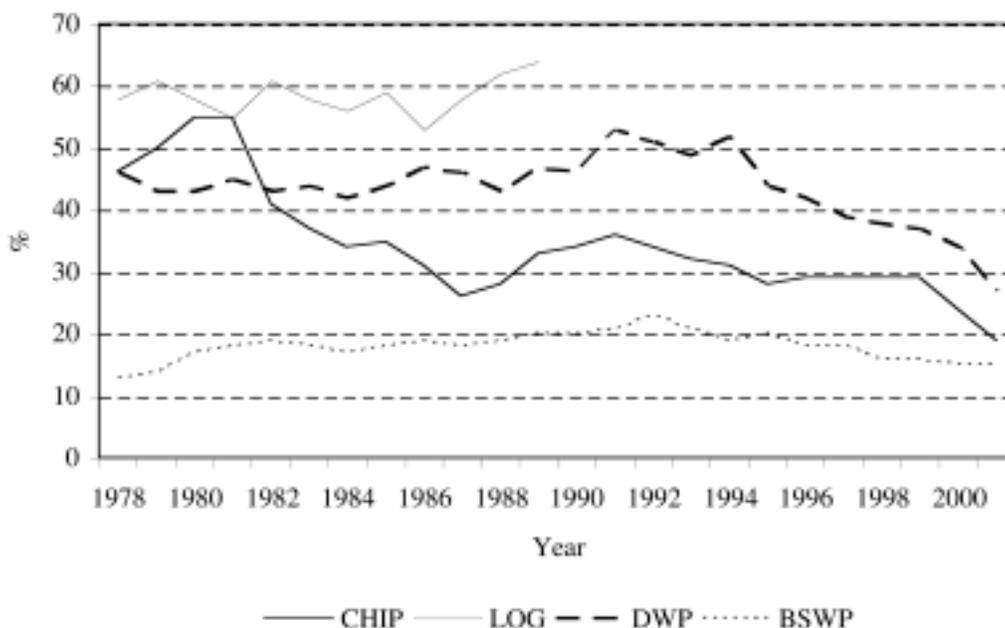


Figure 1. Shares of U.S. export value in total world exports for the selected forest commodities.

Table 1. Results of unit root test for nonstationarity and integration order.

Variable	ADF statistic	
	Level	1 st difference
<i>Q-CHIP</i>	-1.40 (2)	-9.90 (1)
<i>Q-LOG</i>	-1.50 (1)	-11.14 (0)
<i>Q-DWP</i>	0.17 (4)	-7.38 (3)
<i>Q-BSWP</i>	-2.44 (2)	-10.77 (1)
<i>P-CHIP</i>	-1.81 (0)	-10.18 (0)
<i>P-LOG</i>	-2.20 (6)	-8.96 (0)
<i>P-DWP</i>	-1.76 (0)	-9.85 (0)
<i>P-BSWP</i>	-2.01 (4)	-7.48 (3)
<i>G-CHIP</i>	-1.41 (0)	-4.54 (2)
<i>G-LOG</i>	-0.39 (0)	-6.57 (0)
<i>G-DWP</i>	-1.16 (2)	-9.35 (1)
<i>G-BSWP</i>	-1.29 (2)	-8.87 (1)
<i>V</i>	-2.05 (6)	-6.63 (0)

NOTES: (1) The numbers in parentheses are lags decided by the Akaike information criterion. (2) The critical values for ADF test are $\tau(1\%) = -3.50$, $\tau(5\%) = -2.89$, and $\tau(10\%) = -2.58$. The null hypothesis of unit root is rejected for all the series in the 1st difference but failed in the level so all series are nonstationary and integrated of the order one.

The maximum eigenvalue test and trace test are used to verify for common stochastic trends among the series. The null hypothesis is that there are r cointegrating vectors at most, whereas the alternative hypotheses are $r + 1$ and at least $r + 1$ for the λ_{\max} and λ_{trace} statistics, respectively. The results from the Johansen likelihood ratio tests for cointegration are reported in Table 2.

Starting with the λ_{\max} test results, the null hypothesis $r = 0$ (no cointegration) was rejected in favor of $r = 1$ for each of the four forest commodities. The calculated test statistics range from a low of 30.69 for *LOG* to a high of 41.52 for *BSWP*. The critical value at the 5% level is 31.46 for *CHIP* and *DWP*, and 27.07 for *LOG* and *BSWP*. Furthermore, the null hypotheses of $r = 1$, $r = 2$, and $r = 3$ cannot be rejected in favor of the alternative hypotheses of $r = 2$, $r = 3$, and $r = 4$, respectively. These results indicate the presence of one cointegrating relationship for each forest commodity.

For the trace test results, similar conclusions are obtained when the null hypothesis of $r = 0$ is tested against the alternative hypothesis of $r > 1$. Also, the null hypotheses $r \leq 1$, $r \leq 2$, and $r \leq 3$, could not be rejected for all the

commodities. In sum, there is one cointegrating vector for each commodity, suggesting that there is a long-term equilibrium relationship among U.S. exports of forest commodities, relative prices, world purchasing power, and exchange-rate volatility.

The estimated cointegrating vectors are given economic meaning by normalizing with respect to export quantity. The normalized equations are obtained by dividing each cointegrating vector by the negative of the estimated export quantity coefficient. These normalized equations yield estimates of the long-term equilibrium elasticities and are reported in Table 3.

The price (*P*) elasticities have expected negative signs for all commodities except *CHIP*. Only the elasticity for *LOG* is significant at the 5% level. These results suggest that exports respond to price changes. The elasticities for foreign economic activity (*G*) have positive signs for most commodities except *LOG*. They are insignificant at the conventional level, implying a fairly small or weak response of U.S. forest commodities exports to changes in foreign economic activity.

Our elasticity estimates for price and foreign economic activity are a little different from some previous studies in the literature. As noted by Riedel (1988), most estimates of income elasticities in export-demand equations, whether for developed or developing countries or for country aggregates or in individual countries, generally range between 2.0 to 4.0. Arize (1997a, 1997b) and Arize et al. (2000) have consistently reported price elasticities less than one but foreign economic activity elasticities larger than one. This contrast may reflect the difference between export demand equations for a country and for specific commodities.

For exchange-rate volatility (*V*), the results show that the elasticities have negative signs for all commodities, and they are significant at 5% level. The magnitude of the elasticity is smaller than one for three commodities and quite large for *CHIP*. Overall, in the long term, the impact of exchange rate volatility on the exports of forest commodities is negative.

Error-Correction Model and Short-Run Dynamics

The Granger representation theorem (Engle and Granger 1987) proves that if there is a cointegrating relationship among a set of I(1) series a dynamic error-correction representation of the cointegrating vector (*Q*, *P*, *G*, *V*) exists. The

Table 2. Cointegration test results.

	Maximum eigenvalue (λ_{\max})				Trace statistics (λ_{trace})			
	$H_0: r = 0$	$r = 1$	$r = 2$	$r = 3$	$r = 0$	$r \leq 1$	$r \leq 2$	$R \leq 3$
	$H_a: r = 1$	$r = 2$	$r = 3$	$r = 4$	$r > 1$	$r > 2$	$r > 3$	$r > 4$
<i>CHIP</i>	36.24	14.97	11.44	8.84	71.49	35.25	20.27	8.84
<i>LOG</i>	30.69	8.00	4.13	0.17	42.99*	12.31	4.30	0.17
<i>DWP</i>	41.52	22.19	6.05	4.15	73.91	32.39	10.21	4.15
<i>BSWP</i>	40.56	16.80	5.60	1.99	64.94	24.38	7.58	1.99
5% cv (trend)	31.46	25.54	18.96	12.25	62.99	42.44	25.3	12.25
5% cv (no trend)	27.07	20.97	14.07	3.76	47.21	29.68	15.4	3.76

NOTES: (1) Lag length is three for *CHIP*, one for *LOG*, four for *DWP* and *BSWP*. (2) The critical values of Johansen test depend on whether a constant or trend is included in the regression. For *CHIP* and *DWP*, there is a trend in the cointegration regression, so the critical value is different from the other two. * significant at the 10% level.

Table 3. Estimates of the cointegrating relationship.

	Q	P	G	V
<i>CHIP</i>	-1.00	5.41 (2.14)	3.95 (1.71)	-8.51* (-6.17)
<i>LOG</i>	-1.00	-5.59* (-6.50)	-0.29 (-1.68)	-0.66* (-5.00)
<i>DWP</i>	-1.00	-0.48 (-2.18)	0.42 (2.15)	-0.60* (-4.57)
<i>BSWP</i>	-1.00	-0.85 (-1.01)	0.44 (1.85)	-1.00* (-4.54)

NOTES: t ratios are in parentheses and have χ^2 (1) distribution; $\chi^2_{(1, 10\%)} = 2.71$; $\chi^2_{(1, 5\%)} = 3.84$. * significant at the 5% level or better.

error-correction model can be estimated for each forest commodity as follows

$$\Delta Q_t = \alpha_0 + \alpha_1 EC_{t-1} + \sum_{i=1}^4 \beta_i \Delta Q_{t-i} + \sum_{i=0}^4 \gamma_i \Delta P_{t-i} + \sum_{i=0}^4 \delta_i \Delta G_{t-i} + \sum_{i=0}^4 \lambda_i \Delta V_{t-i} + e_t \quad (6)$$

where EC_{t-1} is the lagged error correction term, the residual from the cointegrating regression. Since the variables (Q, P, G, V) have a cointegrating vector, EC_{t-1} is an $I(0)$ process and represents the deviation from equilibrium in period t . The error correction model shows how the system converges to the long-term equilibrium implied by the cointegrating regression. The coefficient α_1 represents

the response of the dependent variable in each period to departures from equilibrium. Its size indicates the speed of adjustment of any disequilibrium towards a long-term equilibrium state.

The methodology used to find this representation follows a “general-to-specific” methodology (Hendry 1987). An error-correction model is first estimated for each commodity with the lag of four. The dimensions of the parameter space are then reduced to a final parsimonious ECM specification by imposing statistically insignificant restrictions sequentially or eliminating insignificant coefficients. Results are summarized in Table 4 along with a battery of tests on statistical adequacy of these ECMs. These tests indicate that the ECM for each commodity is adequate for further analysis.

For two commodities, *LOG* and *BSWP*, the error correction term EC_{t-1} is statistically significant and has the expected negative sign. This coefficient gives a measure of the average speed at which export volume adjusts to a change in equilibrium conditions. The absolute values of the error correction terms indicate that the movement of exports towards eliminating disequilibrium varies across commodities. For *LOG*, about 31% of the adjustment occurs in one quarter, while the figure is 9% for *BSWP*. In other words, the adjustment of export volume to changes in the previous period’s disequilibrium would take slightly more than 3 quarters for *LOG* ($1/0.31 = 3.2$) and 11 quarters for *BSWP* ($1/0.09 = 11$). The error correction terms for *CHIP* and *DWP* have the expected negative signs but are insignificant.

Table 4. Regression results for error-correction models.

	<i>CHIP</i>		<i>LOG</i>		<i>DWP</i>		<i>BSWP</i>	
EC_{t-1}	-0.01	(-1.12)	-0.31*	(-2.46)	-0.10	(-1.23)	-0.09*	(-3.40)
ΔQ_{t-1}	-0.25*	(-2.50)	-0.43*	(-2.91)	-0.57*	(-4.81)	-0.32*	(-3.37)
ΔQ_{t-2}	-0.29*	(-2.79)			-0.28*	(-2.27)	-0.36*	(-3.79)
ΔQ_{t-3}			0.08	(0.42)	-0.24*	(-2.36)	-0.09	(-0.92)
ΔQ_{t-4}			0.16	(0.95)			0.23*	(2.62)
ΔP					-0.57*	(-2.75)		
ΔP_{t-1}			1.38*	(2.10)	-0.39**	(-1.83)		
ΔP_{t-2}	0.13	(0.86)	1.00**	(1.74)	-0.36	(-1.47)	-0.31*	(-2.27)
ΔP_{t-3}			0.77	(1.57)				
ΔP_{t-4}	-0.22	(-1.43)						
ΔG							0.16	(1.56)
ΔG_{t-1}	0.36	(1.54)	-0.41**	(-1.98)			0.20**	(1.85)
ΔG_{t-2}			0.07	(0.35)			0.23*	(2.21)
ΔG_{t-3}					0.25**	(1.65)		
ΔG_{t-4}	0.36**	(1.62)	0.32	(1.49)	-0.27**	(-1.83)		
ΔV			-0.24	(-1.57)	-0.08	(-0.89)		
ΔV_{t-1}	-0.03	(-0.31)					0.11*	(2.22)
ΔV_{t-2}							0.11*	(2.17)
ΔV_{t-3}								
ΔV_{t-4}	0.17*	(2.02)			-0.12	(-1.39)		
R^2	0.23		0.68		0.45		0.45	
Adj. R^2	0.14		0.51		0.37		0.38	
F -statistic	2.68		4.03		5.83		5.95	
DW	2.04		1.91		2.00		1.96	
Norm χ^2 [2]	0.88		0.17		0.21		2.95	
$ARCH F$	0.54		0.48		0.51		0.62	

NOTES: (1) Numbers in parentheses are t -statistics. * significant at the 5% level; ** significant at the 10% level. (2) DW (Durbin Watson statistic) tests first-order residual autocorrelation. Normality χ^2 [2] is the Jarque-Bera test for skewness and excess kurtosis of the residuals. It has a chi-square distribution with 2 degrees of freedom. $ARCH F$ is the F-test for autoregressive conditional heteroskedasticity.

The coefficients on foreign income (G) and relative price (P) in the ECMs show the average speed of export adjustment. Most of these coefficients are significant, indicating that changes in relative price and foreign income have significant short-term effects on exports, as expected. In most cases, the price coefficients are larger than the income coefficients, indicating a faster response of export volume to relative price changes.

Finally, the exchange rate volatility is positive and statistically significant for *CHIP* and *BSWP*, but negative and insignificant for *LOG* and *DWP*. This means that *CHIP* and *BSWP* are responsive to short-run exchange rate volatility, but *LOG* and *DWP* are not. The absolute values range from 0.03 to 0.24. These results indicate that the short-term impact of exchange rate volatility depends on the specific kinds of commodities.

Conclusions

The effects of exchange rate volatility on U.S. exports for four forest commodities (wood chips, softwood logs, dissolving wood pulp, and bleached sulphate wood pulp) are consistent in the long term but mixed in the short term. In the long term, the exchange rate volatility has significant negative effects on the exports of all these forest commodities considered. In the short term, the effect is significantly positive for wood chips and bleached sulphate wood pulp, but negative for softwood logs and dissolving wood pulp. So the impact of exchange rate volatility on U.S. exports is negative in the long term, but short-term dynamics depend on individual commodities. In addition, the speed of the average time lag for adjustment of exports to changes in the determinants of U.S. exports and the short-term effects of exchange-rate volatility varies across individual commodities.

Our analysis for softwood sawlogs and veneer logs (*LOG*) only has data of 10 yr due to data constraints. The exports of logs and the impact of exchange rate volatility may have changed in the 1990s due to the U.S. ban on cutting and exports of logs on federal and state lands in the Pacific Northwest. To the extent that one can generalize from a sample of four forest commodities, it is clear that a stable currency policy enhances U.S. forest commodities exports in the long term, but the short-term impact of exchange rate volatility varies for different forest commodities. Astute producers might be able to benefit by targeting domestic or foreign markets based on their commodities and short-term prospects of exchange rate volatility. As the impact of exchange rate volatility on specific commodities can be much different from that of a country's aggregate export level, further research at commodity levels can provide better guidance for U.S. producers and exporters to gain competitiveness in global market.

Endnotes

- [1] For theoretical studies, see De Grauwe (1988), Franke (1991), Sercu and Vanhulle (1992), and Dellas and Zilberfarb (1993); for empirical studies, see Thursby and Thursby (1987), Koray and Lastrapes (1989), Asseery and Peel (1991), and Chowdhury (1993).
- [2] That is, Q_t , P_t , G_t , and V_t (variables representing export quantity, relative price, world effective purchasing power, and exchange rate volatility, respectively) in this study. See next section.

Literature Cited

- ADAMS, D.M., B.A. MCCARL, AND L. HOMAYOUNFARROKH. 1986. The role of exchange rates in Canadian-United States lumber trade. *For. Sci.* 32(3):973–988.
- ALAVALAPATI, J., W.L. ADAMOWICZ, AND M. LUCKERT. 1997. A cointegration analysis of Canadian wood pulp prices. *Am. J. Agric. Econ.* 79 (August):975–986.
- ANDERSON, M., AND P. GARCIA. 1989. Exchange rate uncertainty and the demand for U.S. soybeans. *Am. J. Agric. Econ.* 71(3):721–729.
- ARIZE, A.C., T. OSANG, AND D.J. SLOTTJE. 2000. Exchange-rate volatility and foreign trade: Evidence from thirteen LDCs. *J. Bus. Econ. Stat.* 18(1):10–17.
- ARIZE, A. 1997a. Foreign trade and exchange-rate risk in the G-7 countries: cointegration and error-correction models. *Rev. Finan. Econ.* 6 (1):95–112.
- ARIZE, A. 1997b. Conditional exchange-rate volatility and the volume of foreign trade: Evidence from seven industrialized countries. *South. Econ. J.* 64:235–254.
- ASSEERY, A., AND D.A. PEEL. 1991. The effects of exchange rate volatility on exports. *Econ. Lett.* 37:173–177.
- BABA, Y., D.F. HENDRY, AND R.M. STARR. 1992. The demand for M1 in the USA, 1960–88. *Rev. Econ. Stud.* 59:25–61.
- CHOWDHURY, A.R. 1993. Does exchange rate volatility depress trade flows? Evidence from error-correction model. *R. Econ. Stat.* 75 (Nov.):700–706.
- COTE, A. 1994. Exchange rate volatility and trade. Bank of Canada Work. Pap. 94-5.
- DE GRAUWE, P. 1988. Exchange rate variability and the slowdown in growth of international Trade. *Internat. Monet. Fund Staff Pap.* 35:63–84.
- DELLAS, H., AND B.Z. ZILBERFARB. 1993. Real exchange rate volatility and international trade: A re-examination of the theory. *South. Econ. J.* 59:641–647.
- ENDERS, W. 1995. *Applied econometric time series*. Wiley, New York. 443 p.
- INGLE, R., AND C.W. GRANGER. 1987. Cointegration and error correction: Representation, estimation and testing. *Econometrica* 55:251–276.
- FRANKE, G. 1991. Exchange rate volatility and international trade. *J. Internat. Money Fin.* 10:292–305.
- FROOT, K.A., AND P. KLEMPERER. 1989. Exchange rate pass-through when market share matters. *Am. Econ. Rev.* 79:637–654.
- HANNINEN, R. 1999. Exchange rate pass-through in Finnish sawnwood exports to the United Kingdom. P. 41–54 in *Modern time-series analysis in forest products trade*, Abildtrup, J., et al. (eds). Kluwer Academic Publishers, Boston, MA. 217 p.
- HASSAN, M.K., AND D.R. TUFTE. 1998. Exchange rate volatility and aggregate export growth in Bangladesh. *Appl. Econ.* 30:189–201.
- HENDRY, D.F. 1987. *Econometric methodology: A personal perspective*. P. 29–48 in *Advances in econometrics*, T. Bewley (ed.). Cambridge University Press, Cambridge, MA. 267 p.

- JOHANSEN, S. 1988. Statistical analysis of cointegrating vectors. *J. Econ. Dynam. Control* 12:231–254.
- JOHANSEN, S. 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. *Econometrica* 59 (November):1551–1580.
- JOHANSEN, S., AND K. JUSELIUS. 1990. Maximum likelihood estimation and inference on cointegration with applications to demand of demand. *Oxford Bull. Econ. Stat.* 52 (May):169–210.
- KENNEDY, P. 1998. A guide to econometrics. Ed. 4. The MIT Press, Cambridge, MA. 468 p.
- KORAY, F., AND W.D. LASTRAPES. 1989. Real exchange rate volatility and U.S. bilateral trade: A VAR approach. *Rev. Econ. Stat.* 71:708–712.
- LEAHY, M.P. 1998. New summary measures of the foreign exchange value of the dollar. *Fed. Reserve Bull.* (Oct.):811–818.
- MCKENZIE, M.D. 1999. The impact of exchange rate volatility on international trade. *J. Econ. Surv.* 13 (Jan.):71–106.
- RIEDEL, J. 1988. The demand for LDC exports of manufactures: Estimates from Hong Kong. *Econ. J.* 98:138–148.
- SERCU, P., AND C. VANHULLE. 1992. Exchange rate volatility, international trade, and the value of exporting firms. *J. Bank. Fin.* 16:155–182.
- SIMS, C.A. 1980. Macroeconomic and reality. *Econometrica* 48:1–48.
- THURSBY, M.C., AND J.G. THURSBY. 1987. Bilateral trade flows, the Linder hypothesis and exchange risk. *Rev. Econ. Stat.* 69: 488–495.
- UUSIVUORI, J., AND J. BUONGIORNO. 1991. Path-through exchange rates on prices of forest products exports from the United States to Europe and Japan. *For. Sci.* 37(3):931–948.

APPENDIX I. Data Construction

Commodities Selection and Time Period

Trade data are from the U.S. Bureau of the Census (USBC) and the Food and Agricultural Organization (FAO). Trade classification systems adopted by both agencies have evolved over time. FAO used the Standard International Trade Classification system (SITC, Rev. 2) from 1978 to 1988 and SITC (Rev. 3) since 1989. USBC used SITC (Rev. 2) from 1978 to 1988 and the Harmonized System since 1989.

Two factors are considered in selecting forest commodities for analysis. First, the selection was constrained by the data availability of other countries' export prices worldwide. The world export prices are reported in the *FAO Yearbook, Forest Products 1978–2001*. To facilitate comparisons between countries, the FAO only reports the trade value and quantities at more aggregated levels. Only those commodities or aggregates that have trade data reported by the FAO can be considered. Second, we intended to select commodities for which the United States is an important exporter in the international market.

Four forest commodities are selected. They are wood chips and particle (*CHIP*), softwood sawlogs and veneer logs (*LOG*), dissolving wood pulp (*DWP*), and bleached sulphate

wood pulp (*BSWP*). The corresponding codes in SITC (Rev. 2) are 246.02 for *CHIP*, 247.10 for *LOG*, 251.60 for *DWP*, and 251.72 for *BSWP*. The corresponding codes in SITC (Rev. 3) are 246.11 and 246.15 for *CHIP*, 251.30 for *DWP*, and 251.51 and 251.52 for *BSWP*. The commodity of *LOG* does not exist in SITC (Rev. 3), so its sample is from 1978 to 1988 only. Other three commodities have the full sample from 1978 to 2001. The United States has been a major player in the international market for all four commodities (Figure 1). For other forest commodities, the average U.S. share of total world export markets has not been more than 10% in the study period.

U.S. Export Prices and Quantities

USBC reports the monthly values and quantities of exports in its monthly publication, *U.S. Exports, Schedule E, Commodity by Country* from 1978 to 1988. For every commodity, the quarterly price is calculated from its monthly value and quantity. From 1989 to 2001, the trade data are abstracted from the interactive online trade database, i.e., *ITC Trade DataWeb*, hosted by the U.S. International Trade Commission. The data are reported in several trade classification systems, including SITC (Rev. 3) and the Harmonized System. In addition, the USBC and the FAO use different units for the four forest commodities selected. The metric units used by the FAO are adopted in this study so USBC data are transformed to the same units.

World Export Prices

For each commodity, the price of world exports is constructed as the export-value-weighted average of the export prices of major exporters in the world market, excluding the United States. Major exporters have a market share of more than 1%. The export price for each country is the ratio of its export value in U.S. dollars to its export quantity. Export values and quantities for each forest commodity by country are from the *FAO Yearbook*, and they are reported on annual basis. The resulted annual world export prices have been converted to quarterly by the quadratic interpolation method, following the method in the literature (Arize 1997a, Arize et al. 2000).

World Income

The world effective purchasing power is measured by the trade-value-weighted average of the real gross domestic product (GDP) of major importers. The series is constructed for each commodity with different weights and country lists. The weight for each commodity is the share of the export value from the United States to the importing countries in the total trade value. The country list for each commodity includes the countries that import over 1% of the U.S. export value. U.S. export value for each commodity by major importing countries is from the same data sources as stated above for U.S. export prices and quantities. The quarterly GDP series in local currency for importing countries are from the *International Financial Statistics* (IFS) published by the International Monetary Fund. GDP data are converted into real value in U.S. dollars with bilateral exchange rates between the United States and each country and the U.S. GDP deflator (IFS).

Exchange Rate of the U.S. Dollar

Since this study involves U.S. exports of forest commodities to over 20 countries, it is more appropriate to use a broad real effective exchange rate index. The BROAD index created by the Federal Reserve Board is adopted (Leahy 1998). The BROAD index is a weighted average of the foreign exchange values of the U.S. dollar against the

currencies of a large group of major U.S. trading partners, including 35 currencies until the beginning of Stage III of the European Economic and Monetary Union in 1999. The index value in 1973 is assumed to be 100. The real version of the BROAD index removes inflation by replacing the nominal bilateral rates with their real counterparts using the consumer price index.