Finnish and Swedish Sawnwood Exports to the UK Market in the European Monetary Union Regime

Jyri Hietala, Riitta H. Hänninen, and Anne Toppinen

Abstract: The business environment for sawnwood producers based in the traditional forest industry countries of the boreal zone has changed dramatically in the last decade. Among the greatest changes is the establishment of the European Monetary Union (EMU). However, the effect of the associated exchange rate changes on international forest industry trade in European markets has not previously been examined. The present study attempts to fill this gap by analyzing the effects of currency movements on Finland's and Sweden's competitiveness in the international sawnwood market. Johansen's multivariate cointegration method was applied for formulation of the separate partial equilibrium model systems for bilateral trade flows from Finland and Sweden to analyze pricing of sawn softwood for the UK market in 1995–2008. Rather than competitive markets, our findings indicate the existence of price discrimination and market power on the part of exporting firms. The results show that Finnish exports have been greatly affected by the currency movements between euros and British pounds in the UK market. Furthermore, the estimated values for exchange rate pass-through coefficients indicate that Finnish exporters have passed on some of the exchange rate changes between these currencies to the UK customers. This result is in line with similar estimations for the period before the EMU. The present estimates show a slightly lesser degree of exchange rate pass-through, possibly suggesting transition toward a more competitive environment in the UK's sawn softwood market. FOR. SCI. 59(4):379–389.

Keywords: exchange rate pass-through, trade, pricing, sawnwood, cointegration analysis

THE BUSINESS ENVIRONMENT FOR SAWNWOOD PRODUCERS in the traditional forest industry countries of the boreal zone has changed in many ways in the last decade or so. Not least among these changes has been the effect of new producers from low-cost countries, further tightening the competition for global market shares, along with the introduction of restrictions to wood availability imposed through forest conservation programs (Hänninen and Kallio 2007) and, more recently, through increased Russian export tariffs on roundwood (Solberg et al. 2010).

However, a change that in some ways is even greater was initiated in European markets by the founding of the European Monetary Union (EMU) in 1999. At that time, many business executives and financial specialists saw mostly positive effects, especially from elimination of the potential for currency speculation and of the exchange rate buffer in intra-EMU trade (Hulkko 2008). However, keen speculation was directed also at whether this could cause severe adaptation problems to, for example, Finnish forest industry firms for which exchange rate devaluations have traditionally had an important role. The EMU was seen as yielding a major benefit for producers operating inside the EMU, thanks to lowered transaction costs caused by the elimination of exchange rate risks. For example, a study by Bun and Klaassen (2002) revealed that the euro has increased intra-EMU trade, with the magnitude of that effect being considerable, a 40% increase in the long-term.

The opinion that exchange rates play a key role in the international trade flows of forest industry products is commonplace. Previous studies concerning forest product exports have directly estimated the elasticity of exports to changes in exchange rates (e.g., Adams et al. 1986, Bolkesjø and Buongiorno 2006), analyzed the negative effects of exchange rate uncertainty on trade volumes (e.g., Sun and Zhang 2003), or reported the use of exchange rate changes to gain price advantages against competitors, especially as a consequence of deliberate currency depreciation (e.g., Uusivuori and Buongiorno 1991, Alavalapati et al. 1997, Hänninen 1998a, Hänninen and Toppinen 1999). However, research results concerning the effects of exchange rate changes after the introduction of the EMU remain absent.[1]

In this study, we try to fill this research gap, analyzing the role of exchange rates in the pricing of forest product exports by applying the concept of the exchange rate pass-through. Whereas previous research indicates that, in the case of Finland, exporters have used exchange rate-induced price changes more fully in their sawnwood exports in comparison with paper commodities (see Hänninen 1998a, and Hänninen and Toppinen 1999), this analysis pertains to the sawnwood trade. We have chosen the UK's sawnwood market as a focus in our analysis, because the United Kingdom is the most important import market in Europe, with Finland and Sweden as two major sawnwood exporters

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and competitors. Accordingly, in addition to that from Finland, we analyze bilateral sawnwood trade flow from Sweden.[2]

From a theoretical angle, we analyze the role of exchange rates in export of forest products by applying the concept of the exchange rate pass-through as a long-run phenomenon. The main inference concerning the exchange rate pass-through phenomenon is related to its impact on price competitiveness of exporters. Much debate has surrounded whether firms pass through some, none, or all of the currency change to export prices denominated in the importer's currency. For a small open economy such as that of Finland or Sweden, the assumption is that firms are price takers such that under perfect competition conditions only profit margins will be affected when exchange rates change.

The research setting is of particular interest currently because the main competitor of the Finnish sawnwood industry, Sweden, has decided to remain outside the EMU and hence serves as a reference point for any effects the loss of an independent monetary policy in Europe has had. In retrospect, exports of Finnish forest products within the EMU region in the 2000s have actually been decreasing despite decreased transaction costs since establishment of the monetary union. As can be detected from Figure 1, after joining the EMU in 1999 Finland lost UK market share as well, while Sweden has been able to increase its share. Another serious consideration has been prompted by the weakening of the krona and strengthening of the euro in the 2000s, which may have raised the price competitiveness of Swedish exporters in euro-based trade. By comparing exchange rate effects between the two exporters, one can gain a more complete picture of the competitive situation encountered in the sawnwood market. Indeed, it is noteworthy that a key issue lies not only in how the value of the euro develops but also in how it develops in relation to other currencies and what kind of pricing strategies are used by Finnish exporters and their competitors outside the monetary union. This consideration has been overlooked in previous studies, in which the final empirical analysis has been based on a *ceteris paribus* condition regarding information of competitors.

Empirically, in this study we apply an econometric method for the separate partial equilibrium model systems for each bilateral trade flow from Finland and Sweden to the United Kingdom. Similar studies in the past have often been based on a mark-up assumption, under which a single-equation method as demonstrated by Engle and Granger (1987) has been commonly applied. In the present study, this conventional method was replaced with a multiequation formulation by Johansen (1988). This allows us to implement the empirical estimation in a system framework, which should yield a more comprehensive view of the phenomenon studied. The majority of the more recent studies have applied Johansen's methodology for modeling the export price formation and trade of forest products (e.g., Sarker 1996, Alavalapati et al. 1997, Hänninen 1998a, Hänninen and Toppinen 1999, Jee and Yu 2001, Sun and Zhang 2003, Hänninen et al. 2007, Nagubadi et al. 2009).

The article is organized as follows. The economic framework and the relevant useful models are presented in the next section, followed by a discussion of data and methods. The remaining sections will then report results of the empirical estimation and draw conclusions based on them.

The Economic Framework and the Models Used

Export Demand

For estimating the exchange rate pass-through, a model system using an export demand equation and a price equation was applied as in, for example, the work of Kongsted (1998) and Hänninen (1998a). The derivation of the export demand assumes a two-stage optimization (see, e.g., Armington 1969).

Because sawnwood is an intermediate product, the export demand equation can be derived from a cost function related to the representative industrial end user of sawnwood in the destination country. First, costs are minimized, in terms of the total expenditure on sawnwood, and, second, this expenditure is allocated optimally between the products from different countries of origin. The following demand relation is obtained for a single exporting country

$$X_i = b_i^{\eta} X_o (P_i/(P_o ER_i))^{-\eta_i} \quad (1)$$

where X_i and P_i are the importing country's demand and the price of the product from export country i. In our case, the United Kingdom represents the importing country, the product is sawnwood, and the exporter is Finland or Sweden. X_o and P_o represent, respectively, the quantity and price of sawnwood from origins other than Finland or Sweden. P_i is expressed in the exporter's home currency, and P_o is the price from the other countries in British pounds (GBP). ER_i is the nominal exchange rate between euro and GBP (EUR/GBP) or between Swedish krona and GBP (SEK/GBP), η_i is the constant price elasticity of substitution, and b_i is a constant.

After logarithmic transformation of the variables in Equation 1, the relation becomes

$$x_i = -\eta_i(p_i - p_o - er_i) + x_o + a_0, \quad (2)$$

where lowercase letters denote the logarithmic values represented by the corresponding uppercase letters in Equation 1. The symbol η_i is the price elasticity of demand for sawnwood, and the constant term is given as a_0. The model assumes that there exists constant elasticity of substitution between sawnwood from different origins, such that market shares are affected only by price relations and not by the size of the market itself (e.g., Armington 1969).

Export Price

The derived export demand equation stated that the final demand decision is based on the relative prices from different countries of origin. We assume that the exporter maximizes profit V_i by taking the competitors' price and supply of sawnwood as given and by setting the price in home currency P_i as a constant mark-up over the unit production costs C_i [$V_i = (P_i - C_i) X_i$].

The first-order conditions for profit maximization imply that the firm equates the marginal revenue from sales to the

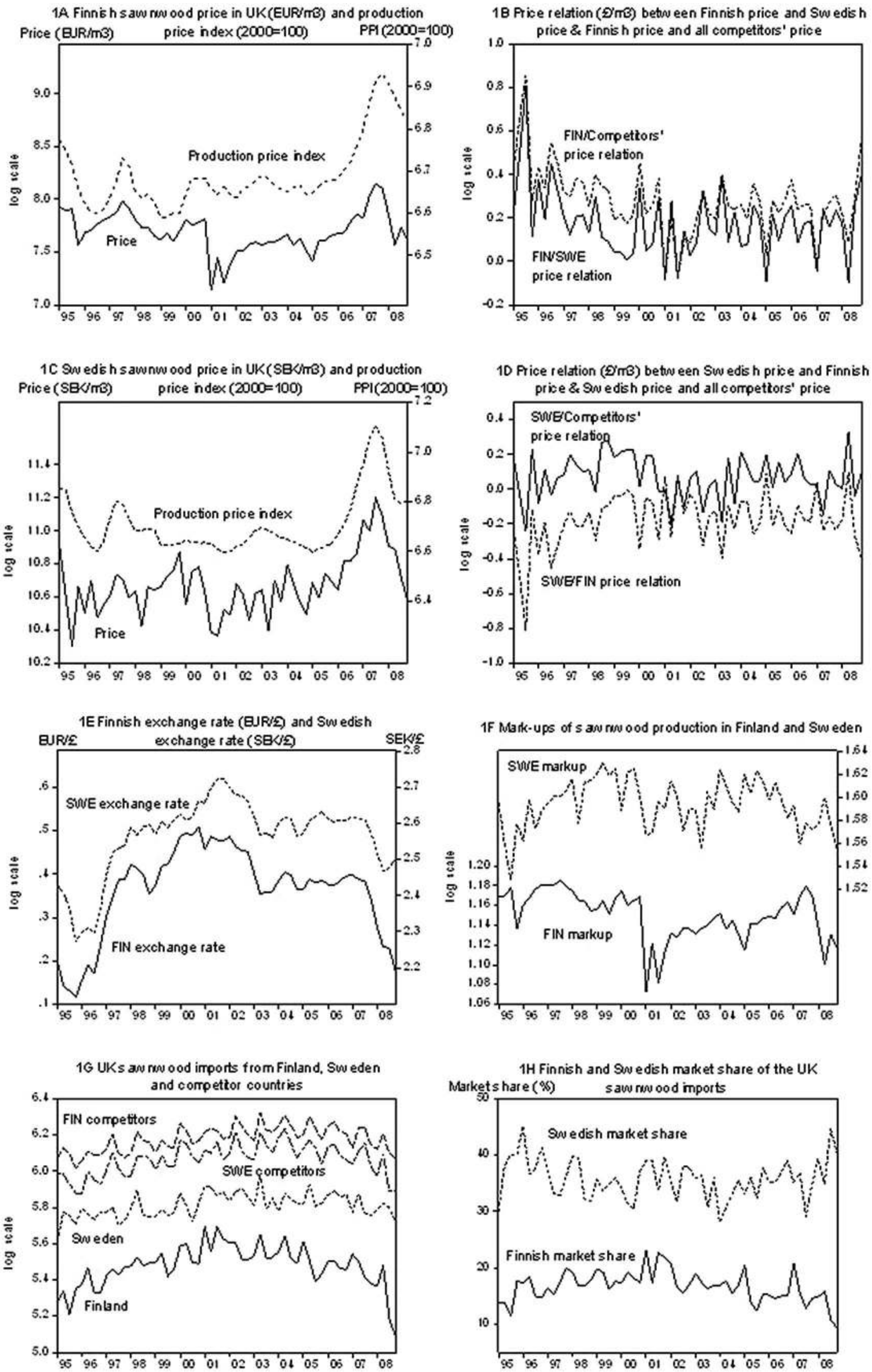

Figure 1. Time series data for UK imports from Finland and Sweden between 1995 (first quarter)–2008 (fourth quarter). PPI, production price index; SWE, Swedish; FIN, Finnish. Sources: The Research Institute of the Finnish Economy database, Eurostat database, the National Bank of Sweden, the Bank of Finland, Statistics Finland database, and Statistics Sweden database.

marginal cost of production [$P_i = C_i\eta/(\eta - 1)$]. Hung et al. (1993), however, showed that by relaxing the constant mark-up condition, one can solve a more general case in which competitors' prices determine the exporter's price. This can be demonstrated by the concept of a variable mark-up. Assuming the mark-up to be variable and to respond to destination-specific market conditions depicts a nonconstant elastic demand (Adolfson 1999). The variable mark-up ϕ is obtained by assuming that the price elasticity of demand η is a function of price competitiveness, $(P_oER_i)/P_i$, in the export market: $\phi = \phi[(P_oER_i)/P_i] = \phi'[(P_oER_i)/P_i]^\theta$, where θ (≥ 0) reflects the relative price elasticity of the mark-up. In the case of a constant mark-up, $\theta = 0$ and $\phi' = \eta/(\eta - 1)$. By substituting the profit-maximizing behavior in the case of a variable mark-up, $P_i = \phi C_i$, together with the notation for variable mark-up ϕ, a relation for the estimated price can be obtained and shown to be

$$p_i = \delta + (1 - \gamma)(er_i + p_o) + \gamma c_i + u \qquad (3)$$

where $\gamma = 1/(1 + \theta)$, $0 < \gamma \leq 1$, $\delta = \ln\phi'/(1 + \theta)$ is a constant, and u is an error term. The price of competing exports, p_o, and exchange rate, er_i, enter the relation with equal coefficients, and, therefore, in estimation the equality restriction is imposed on these coefficients. The export price equation is homogeneous of degree 1 in nominal variables (p_o, er_i, and c_i).

Exchange Rate Pass-Through

The exchange rate pass-through (ERPT) coefficient can be measured as a short- and long-run relationship between changes in exchange rates and export prices. Our emphasis is on the long-run coefficients that depict the steady-state equilibria between currency changes and export prices. In theory, the value of the ERPT coefficient is closely related to the structure of the market, providing information about the degree of competition in the market. The degree of pass-through is a function of the elasticity of demand and supply, and it can be derived as the absolute value of the exchange rate elasticity of export price measured in foreign currency. In the present study, this will be obtained from the price relation (see Equation 3) by first converting the home currency price to foreign currency and then taking a partial derivative with respect to the exchange rate, er_i

$$\text{ERPT} = -\left(\frac{\partial(P_i - er_i)}{\partial er_i}\right) = \gamma, \qquad 0 < \text{ERPT} \leq 1 \qquad (4)$$

where γ measures the impact of exchange rate changes on the foreign currency export price for a given cost (and other explanatory variables) and can therefore be called the pass-through coefficient. The remainder of the exchange rate change, $1 - \gamma$, will be shifted into the domestic currency price by adjusting the mark-up accordingly.

Two polar cases can be distinguished from Equation 4 by following the "small open economy" context, in which nonconstant elasticity of demand is assumed (Swift 2001). An ERPT value of 0 ($\gamma = 0$) would indicate that producers do not possess market power to change the foreign currency price as a result of an exchange rate change. In Equation 3, this would mean that only competitors' prices enter the price equation and changes in the exchange rate are fully absorbed by the variable mark-up. Thus, competition is perfect and the law of one price (LOP) holds. For example, in the case of home currency depreciation (appreciation), the foreign currency export price would remain constant and instead the producer's mark-up would change, increasing (decreasing) profits. The earlier LOP analysis on the UK sawnwood markets for the period 1978−1992 showed no support for the law between Finnish and Swedish sawnwood prices (Hänninen 1998b). On the other hand, the LOP tests indicate integration between the prices of sawlogs and sawnwood in the Nordic countries (e.g., Størdal and Nyrud 2003 and Thørsen 1998). Therefore, Thørsen (1998) concluded that the Nordic countries sell their sawnwood in highly competitive markets, which tends to pressure down and equalize timber prices between countries in the long run. However, more recent results on the LOP are missing.

If, however, the competition is imperfect and producers are able to vary the export price, the ERPT coefficient is between 0 and unity. In the case in which the ERPT value is 1 ($\gamma = 1$), the home currency price stays proportional to production costs, implying that only the foreign currency price had changed as a result of an exchange rate change, keeping the mark-up constant. This could indicate that during depreciations exporters have transferred the exchange rate advantage to benefit customers. In lowering the relative price of the exported product, the purpose has been to gain price competitiveness and possibly increase market share. Where there is currency appreciation, a high ERPT value could, on the contrary, imply that exporters are reluctant to decrease profit margins.

In addition to firms' strategic pricing behavior, exchange rate pass-through estimates are intended to capture the natural change in prices brought about by cost changes of imported inputs as well. If these are significant, the ERPT coefficient will tend to be underestimated (Athukorala and Menon 1994). In the present study, input cost effects can be assumed to be minor, because the forest industries in Finland and Sweden use mainly domestic inputs.

Empirical Models

The estimable model systems formed for Finland and Sweden include two equations as follows

Finnish model system

$$\textit{Demand}: \quad x_{\text{FU}} = -\eta(p_{\text{FU}} - p_{\text{CU}} - er_{\text{FU}}) + x_{\text{CU}} + a_0 + \varepsilon \qquad (5)$$

$$\textit{Price}: \quad p_{\text{FU}} = \delta + (1 - \gamma)(er_{\text{FU}} + p_{\text{CU}}) + \gamma c_{\text{F}} + u \qquad (6)$$

Swedish model system

$$\textit{Demand}: \quad x_{\text{SU}} = -\eta(p_{\text{SU}} - p_{\text{CU}} - er_{\text{SU}}) + x_{\text{CU}} + a_0 + \varepsilon \qquad (7)$$

$$\textit{Price}: \quad p_{\text{SU}} = \delta + (1 - \gamma)(er_{\text{SU}} + p_{\text{CU}}) + \gamma c_{\text{S}} + u \qquad (8)$$

where x_{FU} and p_{FU} are the Finnish and x_{SU} and p_{SU} are the Swedish sawnwood export quantities (m^3) and nominal home currency unit prices (EUR/m^3 and SEK/m^3) in the United Kingdom and where x_{CU} represents competitors' quantities exported to the United Kingdom and p_{CU} the weighted average foreign currency price (GBP/m^3) of the competitors' products. The nominal exchange rates between the euro and pound sterling (EUR/GBP) and the krona and pound sterling (SEK/GBP) are given as er_{FU} and er_{SU}. c_F and c_S correspond to the Finnish and Swedish sawnwood production cost, respectively. Constant terms are presented via a_0 and δ, and disturbance terms, to catch the effects of all other factors, are given with ε and u.

Data

The trade data of the present study consist of the coniferous sawnwood quantities and unit prices from Finland and Sweden and the rest of the countries for the UK market. The study uses quarterly data, with 56 observations from the beginning of 1995 to the end of 2008. In the empirical estimation, the time span was extended to account for a short pre-EMU period to provide more observations. However, this is not conceived as a major problem in validation of the results obtained, because the "additional period" is rather short and, in any case, should be fairly well in line with the post-EMU period. The decisions already taken to join the EMU should be reflected in the market agents' expectations on the Finnish currency, and hence they should act accordingly. Indeed, the Finnish mark (FIM) had been added to the European Exchange Rate Mechanism already in 1996 to reduce exchange rate variability and achieve monetary stability in preparation for the EMU.[3]

The data on the quantities and values of coniferous sawnwood were gathered from import statistics for the United Kingdom (Eurostat database; European Commission 2010). The product code used in Combined Nomenclature classification is 4407 10, which comprises 14 subproducts in total. The unit values were calculated by dividing total import values (Cost Insurance Freight, CIF, figures in EUR) by the corresponding quantities (m^3). The values were not deflated, because relative prices were used in the analysis. The unit value for countries other than Finland and Sweden (P_o) was calculated for each data point by taking a weighted average for the rest of the importers to the United Kingdom. In the UK sawnwood market during the period under study, Finland's and Sweden's main competitors were Latvia and Russia. To obtain the home currency prices for Finland and Sweden, the respective import unit values were converted, with the average nominal exchange rates used accordingly. The exchange rates were obtained from databases of The Research Institute of the Finnish Economy (2010) and the Swedish National Bank, the Riksbank (2010). Production costs of sawnwood (C_i) were described by production price indices for the wood industry (2000 = 100), which were obtained for Finland from the database of Statistics Finland (2010) and for Sweden from the database of Statistics Sweden (2010). The graphs in Figure 1 describe the data.

Estimation Method

The present study applies Johansen's (1988, 1995) cointegration method. Model systems for Finland and Sweden (5 and 6 and 7 and 8) are estimated separately. Six variables (m = 6) are included in each of the systems. Analysis starts with estimation of a m-dimensional unrestricted vector autoregressive (VAR) process of order k for each of the two model systems. This can be formulated as

$$\Delta X_t = + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-k} + \mu + \Phi D_t + \varepsilon_t \qquad (9)$$

It is assumed that the error term $\varepsilon_t \sim$IID $(0, \Omega)$. To obtain a well-behaving error term, we introduce a sufficient number of lags, k. $\tilde{\Gamma}_i, \ldots, \tilde{\Gamma}_{k-1}$ and $\Pi = -\mathrm{I} + \Pi_1 + \ldots + \Pi_k$ are coefficient matrices. The deterministic terms are expressed by D_t containing, for example, a constant, seasonal dummies, and intervention dummies. $\Pi = \alpha\beta'$ is the matrix of long-run coefficients of the lagged levels, where β is a matrix of cointegration vectors and α is a matrix of loadings. If some $X_t I(1)$ series form a linear combination of which is stationary $I(0)$, the components of it are said to be cointegrated. Johansen's (1988) trace test and the maximum eigenvalue tests are applied for testing the rank of the matrix Π. In the case in which the variables of each of the model systems are cointegrated, it is possible to identify cointegration vectors, the coefficients of which describe the long-run equilibrium relations (model systems 5 and 6 and 7 and 8). Multiple cointegrating vectors exist in the cases in which $1 < \Pi < n$, where the number of variables in each model system is denoted by n.

After determining the number of cointegrating vectors, one can identify them, conditional on r, by imposing linear restrictions on the matrix of cointegration vectors, β, and the loadings, α. The tests are all likelihood ratio tests, in which the model systems are estimated with and without the restrictions on α and β (Helles et al. 1999). In addition to purely identifying restrictions, it is often necessary to test for restrictions that are structural, i.e., for which the hypotheses are not dependent on any normalization of the parameter β. Different forms of restrictions have been discussed in more detail by authors such as Johansen (1988, 1995) and Johansen and Juselius (1990, 1992) and are presented next as applied in the present study.

The cointegration relations are to satisfy the restriction $\beta = H\varphi$ or ($\Pi = \alpha\varphi$'H'), where H is a ($m \times s$) matrix and defines the linear restrictions, s, determined by φ, on parameter β. The same linear restrictions are formulated for all of the cointegrating relations r where $r \leq s \leq m$. The alternative hypothesis for the above restrictions is the hypothesis $\Pi = \alpha\beta'$, where Π is unrestricted. The test statistic is asymptotically distributed as χ^2 with $(m - s)r$ degrees of freedom. In the present study, this hypothesis was applied to test for price homogeneity by restricting the export price, p_i, and other nominal variables accordingly. Another example is given by the overidentification of the price and demand relations by excluding irrelevant variables from the distinctive cointegration vectors. Last, the mark-up pricing assumption was tested in the price relation by its restriction to either unity or zero, depending on the magnitude of the unrestricted value obtained.

Demonstrated by Johansen (1992b) and Johansen and

Juselius (1994), a similar formulation is possible for identifying short-run dynamics by setting hypotheses on the loadings, $\alpha = A\psi$ or ($\Pi = A\psi\beta'$). This is equivalent for testing of weak exogeneity with respect to the long-run parameters. In the present study, all variables were tested for weak exogeneity, where α_{ij} measures the weight by which each cointegrating vector, β_j, enters each of the ith equations in the system and weak exogeneity implies that all the elements of the lth row of α are 0. Thus, the linear restrictions on αs can be seen more as conditioning, whereas restrictions on βs imply a transformation of the process.

Results

Estimation Results for the Model Systems

The empirical analysis begins with testing of the time series for stationarity. The augmented Dickey-Fuller unit root tests indicated that the variables of the two model systems are of $I(1)$ or $I(0)$, implying that the Johansen methodology is applicable (Table 1). However, because stationary variables can have an effect on the number of cointegrating relationships found by the Johansen cointegration test, according to the unit root test results $r \geq 2$ could be found in the system for Finland. The system for Sweden does not have stationary variables in levels outside the already assumed cointegrating vectors, so $r = 2$ is expected.

An unrestricted VAR model with six equations (dimension $m = 6$) as shown by Equation 9 was estimated for Finland and Sweden for 1995–2008. Impulse dummy variables, D_t, to take into account unexplained price movements during the four quarters of 2001, were included as exogenous variables in both model systems. The lag length, $k = 1$, for the Finnish and Swedish model systems was chosen on the basis of the Schwarz and Hannan-Quinn information criteria.

The cointegration rank of the Finnish and Swedish data was tested by means of a specification allowing for a linear trend in X_t, while the cointegrating relations $\beta' X_t$ are stationary or a model that allows for deterministic trends in both X_t and the cointegrating relations. The former means adding a constant in the model, and the latter adds a constant and a trend. Following the suggestion of Johansen (1992a), the specification was chosen in accordance with the Pantula principle (see Pantula 1989), by starting from the most restricted model and proceeding step by step to a more unrestricted one. In addition, a simple inspection of the data series was conducted.

Results regarding the cointegration estimation of the VAR(1) models indicate that there may be more than two cointegrating vectors, $r \geq 2$, in the model for Finland at the 5% level (Table 2). Moreover, the trace test and the maximum eigenvalue test give somewhat conflicting results for the data of the Finnish model, which may be due to the fairly small sample sizes used in the study and the resulting low power of the distinctive tests. Kongsted (1998) emphasized also that the asymptotic distributions of the test statistics are only approximations, because they do not allow for the inclusion of exogenous D_t terms in the VAR model. This may be reflected in the contradictory test results as well. In the present study, $r = 2$ would be consistent with the theoretical presumption that the export demand and domestic price relations are formed simultaneously and equilibrate in the long run. However, because the cointegration rank test is affected by stationary variables and by included deterministic terms, cases in which $r > 2$ were also investigated as indicated by the rank test results. The inclusion of deterministic components was based on the Pantula principle and the significance of the added drift term as well as on inspection of the stationarity of the cointegrating relations.

According to the results, the data for Finland are trend-stationary: the drift term eliminates trends from the cointegration relations, and, in addition, the trend term is highly significant in the vector relations, increasing the likelihood of the whole system. According to the Pantula principle, neither of the model specifications could be rejected for $r \leq 2$ by the trace test. Following Enders (2004), the maximum eigenvalue test was finally chosen for use to pin down the number of cointegration relations. Therefore, because a model comprising an intercept and a trend term was used for further modeling, two cointegrating vectors ($r = 2$) cannot be rejected in the system at the 5% level. On the other hand, the data for Sweden did not show any signs of trend-stationarity, and according to the Pantula principle $r \leq 2$ could not be rejected for the first time when only an intercept was included in the model specification. The presumption of two cointegrating vectors ($r = 2$) thus cannot be rejected for the Swedish model either. The resulting normalized maximum likelihood estimates (MLEs), β_j, and their corresponding weights, α_j, obtained from the cointegration estimations are available from the authors by request.

Of the six eigenvectors in both systems, the first two relations were most highly correlated with the stationary part of the process Δx_t and were thus normalized on export demand, x_i, and home currency price, p_i. These are to be

Table 1. Augmented Dickey-Fuller unit root test results: estimation samples 1995, first quarter–2008, fourth quarter.

Variables	Level (C)	First difference (C)	Level (CT)	First difference (CT)
Finland				
x_{FU}	−1.42	−7.46**	0.30	−8.03**
p_{FU}	−3.22*	−9.58**	−3.18	−9.50**
er_{FU}	−1.86	−4.93*	−0.08	−5.95**
c_F	−2.41	−4.32**	−3.56*	−4.34**
p_{CU}	−2.44	−7.94**	−2.4	−7.82**
x_{CU}	−1.70	−9.87**	0.43	−10.85**
Sweden				
x_{SU}	−2.15	−8.61**	−1.36	−8.94**
p_{SU}	−3.84*	−10.75**	−4.65*	−10.62**
er_{SU}	−2.01	−5.11**	−0.83	−5.24**
c_S	−1.70	−5.42**	−2.08	−5.35**
p_{CU}	−2.27	−4.12**	−2.10	−4.09*
x_{CU}	−1.67	−7.79**	0.40	−8.78**

Critical values are from Dickey and Fuller (1979). Testing was performed using the Schwarz information criterion with maximum five lags and a constant (C) and a constant plus a trend (CT).
* Rejection of the null hypothesis of nonstationarity at the 5% level.
** Rejection of the null hypothesis of nonstationarity at the 1% level.

Table 2. Cointegration rank tests of unrestricted VAR models under $m = 6$ and $k = 1$: estimation samples 1995, first quarter–2008, fourth quarter.

	Linear					
	Intercept (C), no trend			Intercept (C), trend		
Null hypothesis $H_0 = r \leq i$	Eigenvalue λ_i	Trace test probability	Maximum eigenvalue probability	Eigenvalue λ_i	Trace test probability	Maximum eigenvalue probability
Finland						
None	0.53	0.00*	0.03*	0.69	0.00*	0.00*
At most 1	0.40	0.02*	0.23	0.51	0.00*	0.04*
At most 2	0.32	0.04*	0.24	0.39	0.03*	0.16
At most 3	0.26	0.11	0.18	0.31	0.10	0.22
At most 4	0.15	0.29	0.26	0.19	0.28	0.41
At most 5	0.01	0.47	0.47	0.12	0.35	0.35
Sweden						
None	0.65	0.00*	0.00*	0.67	0.00*	0.00*
At most 1	0.48	0.00*	0.03*	0.58	0.01*	0.00*
At most 2	0.31	0.07	0.32	0.33	0.37	0.50
At most 3	0.22	0.13	0.35	0.24	0.55	0.65
At most 4	0.14	0.16	0.31	0.17	0.63	0.58
At most 5	0.05	0.07	0.08	0.07	0.74	0.74

C, constant.
* Rejection of the hypothesis at 5% level.

further overidentified by imposing restrictions driven from the economic framework of the study. The α_{ij}s of the estimated unrestricted MLE model systems were fairly low in the model system for Finland, implying low adjustment, whereas they are larger in the system for Sweden, implying a rapid adjustment process. The α_{ij}s represent the weights with which the error correction terms enter each equation, indicating the speed of adjustment toward the estimated equilibrium state.

Overall, after an inspection of graphs representing the cointegration relations, the vectors seem to be fairly stationary, indicating that the VARs have been reduced to $I(0)$ space. The test results related to the error term, ε_t, are reported in Table 3. The F-form of the Lagrange multiplier test and the White (1980) heteroscedasticity test indicate that the null hypotheses of no autocorrelation and of homoscedasticity could not be rejected at the 5% level for either of the equation systems. The Jarque-Bera normality tests (Jarque and Bera 1980) via Cholesky (JB_{CHO}) and Urzua (JB_{URZ}) factorizations did not reject the null hypothesis of multivariate normality. Because no misspecification test was significant at the conventional 5% level, there is a strong indication that the systems represent the data satisfactorily.

Exchange Rate Pass-Through

The next step is to examine the unrestricted cointegration vectors obtained from the MLE estimation in view of restrictions imposed from economic theory. This is done to identify the systems comprising export demand and price relations for Finnish and Swedish sawnwood in the UK market (model systems 5 and 6 and 7 and 8). Estimable models will measure the total pass-through, i.e., the entire effect an exchange rate change causes, working through every interaction of the price determination. Another possibility is to estimate partial pass-through, the effect an exchange rate change has on the price-setting relation alone, excluding causation of other variables and long-run relations (Adolfson 1999).

Valid conditioning of the model requires that a variable be weakly exogenous for the long-run parameter, β (Kongsted 1998). This implies that the variable does not adjust to deviations from the cointegrating relations. On the one hand, according to demand-supply equilibrium, it is expected, a priori, that at least the prices of exports and export quantities adjust. On the other hand, exchange rates are not expected to respond to changes in export quantities and domestic prices, because exporters cannot affect exchange rates with their own strategies. Indeed, the null of weak exogeneity for exchange rates could not be rejected at the 5% level in both model systems, implying a weak feedback to both currencies. This result was as expected, because exchange rates are supplied to the exporters of both countries: in Finland by the monetary policy of the European Central Bank and in Sweden by the Swedish National Bank, the Riksbank.

Table 3. Misspecification tests on model systems under $r = 2$: estimation samples 1995, first quarter–2008, fourth quarter.

	Autocorrelation (LM)	Heteroscedasticity (white)	Normality	
			(JB_{CHO})	(JB_{URZ})
Finland	34.74 (0.52)	359.64 (0.45)	14.17 (0.28)	112.60 (1.00)
Sweden	45.70 (0.12)	378.96 (0.20)	19.22 (0.08)	126.52 (0.99)

Marginal significance level in parentheses. LM, Lagrange multiplier.

Table 4. LR tests of structural hypotheses on model systems under $r = 2$: estimation samples 1995, first quarter–2008, fourth quarter.

Finland					Sweden				
Variables/ LR test	(I) Long-run homogeneity		(II) Mark-up pricing		Variables/ LR test	(I) Long-run homogeneity		(II) Mark-up pricing	
	β_{i1}	β_{i2}	β_{i1}	β_{i2}		β_{i1}	β_{i2}	β_{i1}	β_{i2}
x_{FU}	1.00	0.00	1.00	0.00	x_{SU}	1.00	0.00	1.00	0.00
p_{FU}	2.38	1.00	2.14	1.00	p_{SU}	0.93	1.00	0.23	1.00
er_{FU}	−2.38	−0.32	−2.14	0.00	er_{SU}	−0.93	−0.82	−0.23	−1.00
c_F	0.00	−0.68	0.00	−1.00	c_S	0.00	−0.18	0.00	0.00
p_{CU}	−2.38	−0.32	−2.14	0.00	p_{CU}	−0.93	−0.82	−0.23	−1.00
x_{CU}	−1.51	0.00	−1.51	0.00	x_{CU}	−0.46	0.00	−0.48	0.00
LR	$\chi^2 = 13.13$		$\chi^2 = 13.71$		LR	$\chi^2 = 22.85$*		$\chi^2 = 23.24$*	
Probability	0.07		0.09		Probability	0.00		0.00	

β_{i1} and β_{i2} correspond to export demand and price relations respectively. LR, likelihood ratio.
* Rejection of the restricted model system at 5% level.

The next step was to identify the model systems by restricting the long-run coefficients accordingly. The tests for the structural hypothesis on the model systems are presented in Table 4. Constant terms were included in the estimation but have been left out from the reported results to save space. The same applies to added dummy variables and trend factors. From the unrestricted cointegration vectors obtained from the MLE estimation, the first relation in the Finnish and Swedish model systems was identified as the demand equation, by excluding production costs, c_i, from the relation ($\beta_{41} = 0$) and by setting the coefficients of competing exports, x_o, to -1 ($\beta_{61} = -1$). The latter restriction was, however, relaxed from both of the models, because the coefficients for competing exports (x_o) in the unrestricted models deviated quite a bit from the proposed value. The implication is that the constant market share assumption does not hold (Figure 1). The second cointegrating vector, β_2, in both model systems was identified as the price equation through exclusion of the variables representing export quantities, x_i and x_o, from the relation ($\beta_{12} = 0$ and $\beta_{62} = 0$).

As was pointed out in the discussion of the economic framework of the study, the theoretical model suggests long-run relations for both export quantity and price to be homogeneous in the nominal variables. These homogeneity restrictions are presented as the first testable long-run structure (I). This means that the coefficients of p_o and er_i should be equal in the two relations ($\beta_{5j} = \beta_{3j}$) and the coefficient of c_i should equal the difference between the coefficients of p_i and er_i in the price vector [$\beta_{42} = -(\beta_{22} + \beta_{32})$] (Hänninen 1998a). The former restriction implies that relative prices determine export demand. In addition, the price of competing products and the exchange rate are anticipated to affect exported quantities as well as the price setting with the same magnitude in the long run. The latter restriction indicates that marginal cost and exchange rate changes would have equal effects on the price measured in local currency (Adolfson 1999).

The results show that the long-run homogeneity restriction was not rejected in the model system for Finland, but it was rejected in the system for Sweden. The latter result is rather peculiar, because the coefficients in the unrestricted model for Sweden indicated only minimal deviation from the homogeneity assumption. Athukorala and Menon (1995) have, however, noted that in practice it is common that these coefficient restrictions are not valid. They point out that exchange rates tend to be more volatile than both world price and production costs are and that firms therefore may be more prone to absorb exchange rate changes into their profit margins. In addition, Bache (2002) emphasized that in the short run, exchange rate changes are often seen as temporary and exporters accordingly will be willing to absorb them in their mark-ups.

The restrictions for long-run homogeneity structure were not rejected with the model for Finland. This result is in accordance with earlier findings regarding ERPT estimates for Finnish sawnwood exports (e.g., Hänninen 1998a). The price relation is again found to resemble more of a mark-up pricing strategy for which $\gamma = 1 =$ ERPT as the coefficient of production costs, c_F, approaches unity (-0.68). In addition, the relatively large exchange rate elasticity of export demand (2.38) implies a large effect of exchange rate on Finnish sawnwood exports to the United Kingdom. The mark-up assumption (II) was tested by restricting the price relation accordingly ($\beta_{42} = -1$). The probability of this structure increased slightly, so that the equation system could now be nonrejected by a larger margin. The effects of the EMU period on the pass-through can be implicitly evaluated by comparing the coefficient obtained by the present study, -0.68, with that of Hänninen (1998a), -0.92. The lower coefficient for our data indicates a change toward a more competitive environment for sawnwood producers in the EMU period than before the EMU period.

The mark-up pricing structure was tested for the system for Sweden by restricting the pass-through coefficient to zero ($\beta_{42} = 0$), as the coefficient of c_S resembled a pricing strategy (-0.18) opposite that of Finnish exporters. In addition, the fairly small magnitude of the exchange rate elasticity (0.93) implies less effect of the krona value on Swedish sawnwood exports to the United Kingdom than of the euro value on Finnish sawnwood exports. The likelihood in this model lowered slightly, thus still indicating a clear rejection of the overall model structure. The result further indicates that, although Swedish exports of sawnwood to the United Kingdom have been less sensitive to exchange rate changes than have Finnish sawnwood exports, the hypothesis of the effect on prices and exported quantities cannot be completely rejected.

Moreover, coming to far-reaching conclusions as to the pricing strategy of Swedish sawnwood exporters is difficult and rather unreliable. The exceptionally high values of some loadings in the system for Sweden imply misspecification problems with the model system. Because the Johansen procedure does not perform well for small samples, these problems may have arisen also from the use of a heavy estimation system with a fairly large number of variables in proportion to the number of total observations. This apparent drawback of the Johansen multiequation method has been recognized by several authors (e.g., Adolfson 1999).[4]

Discussion and Conclusions

One of the major changes in the business environment of forest product industries in Europe was set off by the founding of the EMU in 1999. The aim of this study was an empirical evaluation of the effects of EMU participation from the standpoint of pricing strategies of the Finnish and the Swedish sawnwood industry. Our fundamental emphasis was on the long-run effects, for purposes of illuminating the strategic pricing induced by exchange rate fluctuations. Consequently, the system of models was based on the theoretical presumption of a small open economy, in which exporters are assumed to be price takers and the relative prices of competitors determine the quantities exported from each country of origin. The empirical results for 1995–2008 are, however, somewhat in contrast with this assumption, appearing more consistent with price discrimination and the existence of market power of exporting firms.

The estimated ERPT coefficients indicate that Finnish exports have been affected to a great extent by currency movements in the UK market. Depreciation experienced in the first half of the observation period has stimulated export demand, whereas appreciation in the second half has, in turn, dampened imports to the United Kingdom from Finland. This result is broadly in line with previous estimations of a similar scope (e.g., Hänninen 1998a, Hänninen and Toppinen 1999). However, the ERPT estimated from the present data (-0.68), was lower than that estimated from the data before 1995 (-0.96) (Hänninen 1998a). This result implies the general increase of competition in the European sawnwood markets. In the economic fluctuations of late 1980s and early 1990s, Finnish exchange rate policy was supportive to the competition ability of the Finnish export industry, and there were several realignments of the Finnish currency (the markka). This could have generated a greater pass-through rate. Monetary policy in the EMU is managed by the European Central Bank, and the euro exchange rate has been more stable, which, in effect, could have led to a more conservative pricing strategy of exporting firms. For example, Devereux et al. (2004) concluded in their study that the more stable the monetary policy is, the lower the relative pass-through rate too will be.[5]

In contrast, the pricing strategy exploited by Swedish exporters has been somewhat the opposite of that of Finnish exporters, even though the development of the respective currencies against the British pound has been fairly parallel for the period since 1995. This has meant both a more stable Swedish sawnwood price for the customers in the United Kingdom and a more stable export demand faced by Swedish exporters. Our results further suggest that depreciation of the krona against the pound sterling has not negatively affected Finnish exports to the United Kingdom, at least not to the fullest possible magnitude. Still, higher profit margins achieved by Swedish producers appear to have been an important consideration behind some company-level decisions to shift production capacity between Finland and Sweden. One reason for the higher mark-up for Sweden may be the lower total costs of transportation from Sweden to the United Kingdom than from Finland. Finland is situated at a longer distance from the UK markets than Sweden.

In conclusion, the key implication of this study is that despite the realization of the third stage of the EMU in 1999, the decision of the Finnish government to join the EMU seems not to have been a major determinant of the competitiveness of Finnish sawnwood producers in international markets. Other changes in market conditions, e.g., through the overall decrease of the UK's sawnwood imports, large sawmill capacity investments in Western Europe and increased supply from Eastern European countries to EMU markets, have changed the situation. The growing importance of African and Asian markets for Finnish sawnwood exports have exposed Finnish trade more to exchange rate changes between the euro and the other currencies, such as US dollars and yen. In addition, increased supply from Eastern European and Asian countries and recent worldwide dampening of housing demand as a result of the global financial crisis emerging in 2007–2008 and worsening thereafter as well as technological advances could be considered equally important factors affecting the development of Finnish sawnwood exports.

The results of a recent firm-level study (Lähtinen et al. 2008) have indicated the importance of sawnwood industry resources as a source of competitive advantage The competitive advantage of the Finnish sawmilling industry is mainly driven through resources related to personnel, managerial, raw material, and financial factors. In addition, Toivonen et al. (2005) emphasized the importance of sawnwood product quality, in particular, the detection of better intangible product components through higher quality in supplier service, logistics, and other dimensions of the intangible product offering. However, companies appear to be focusing less on areas highlighted by these research findings than on applying the general perception of the forest industry in the Nordic countries; that is, they currently seem more focused on protecting the traditional cost-based determinants of long-run competitiveness, e.g., overall improved production efficiency and good availability of inexpensive energy, wood raw material, and other input factors.

Today's forest-industry companies are multinational corporations whose production capacity spans national borders so that internally production can be and has been shifted to where it is most profitable. For example, Gron and Swenson (1996) have suggested that firms' export prices are unlikely to match exchange rate fluctuations one to one under these circumstances, which could affect the generalizability of the estimation results. An interesting extension to the present study would then be to research pricing strategies on the company level, were such data available. In addition, this would enable comparisons between companies with differing size and scope of foreign trade. In view of the restrictions concerning data used in the present study, some results can be considered tentative. In forthcoming analysis of the international sawnwood trade, other European markets should also be included and the observation period should be lengthened to account for the sole EMU era. Further research would be needed similarly for analyzing forest product categories in greater depth, as well as examining possible asymmetries in pricing strategies (e.g., von Cramon-Taubadel 1998). In addition, from a methodological perspective, studying short-run dynamics of exchange rate changes could provide further insight into export price determination, for example, through consideration of the choice of invoicing currency.

Endnotes

1. Even recently, it has been discussed that one important reason for the growth in exports of Swedish sawnwood during the economic recession after 2007 has been the weakness of the Swedish krona against the US dollar and the British pound (UNECE, Timber Committee 2009).
2. Because the Swedish currency, kronor, has been traditionally used as a pricing currency in the sawnwood trade in Europe, it is of interest in the present study to focus particularly on the sawnwood trade.
3. In addition, Mayes and Virén (2009) suggest that a change in EMU participants' behavior occurred in approximately 1996, when member states were trying to converge under stage 2 of the EMU, with the trend having continued thereafter.
4. To further examine Swedish ERPT estimates, a simpler model comprising only the price relation was tested also; a model similar to that of Hänninen and Toppinen (1999) was applied. The estimation results from the single-equation models resemble those obtained from the multiequation models: a long-run exchange rate pass-through coefficient of 0.26 is obtained for the sawnwood price from Sweden. Hence, the Swedish price relation does not suggest a mark-up pricing strategy, and the exporter's exchange rate changes have seemingly had a much larger effect on price determination in domestic currency. However, some caution has to be advised for not reading too much into the results, because the VAR structure did not pass some of the diagnostic tests on residuals.
5. To study the matter more comprehensively, the Finnish model system was reestimated by including step dummies to separate the pre- and post-EMU periods. In this case, the ERPT obtained the value of 0.53, which would imply a lower effect of exchange rate changes on export price during the EMU era and therefore possibly transition toward a more competitive UK sawnwood market.

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