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Demand for Finnish exports of forest products:

Econometric analyses using time series data

Riitta Hänninen

HELSINGIN TUTKIMUSKESKUS – HELSINKI RESEARCH CENTRE

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Econometric analyses using time series data**

Riitta Hänninen

Doctoral dissertation

To be presented, by the permission of the Faculty of Agriculture and Forestry of the University of Helsinki, for public examination in Auditorium XII of the main building of the University, Unioninkatu 34, on November 13, 1998 at 12 o'clock noon.

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ABSTRACT

The study analyses Finland's and its main competitor countries' export price formation in British (UK) and German (BDR) forest product markets using time series data. The results of the study are useful in modeling forest product trade flows. They also yield information for policy makers and industrial agents, for example, in the adjustment to the third stage of European Economic and Monetary Union (EMU). Johansen's cointegration method, thusfar infrequently applied to the Finnish forest sector, is used in the estimation.

Export price formation is examined in three ways. First, price substitution between Finland's, and its competitor's sawnwood is analysed in the UK market. Second, the law of one price (LOP) related to arbitrage is tested for prices of sawnwood and newsprint originating from Finland, and its competitor countries. Finally, the effects of exchange rate changes on Finnish sawnwood, newsprint and chemical pulp export prices are analysed in both destination countries.

The estimation results indicated imperfect competition in the export markets. The price elasticities of substitution between Finnish, Swedish and Canadian sawnwoods were relatively low and the LOP was rejected in the UK. The LOP was also rejected between the main newsprint supplying countries in German and British markets except between Sweden and Canada in the UK. The resulting exchange rate pass-through (PT) elasticities of Finnish export prices indicate the importance of exchange rates on exports of forest products. Higher PT for sawnwood than for paper suggests that Finnish sawnwood exporters have made use of depreciations and devaluations of the Finnish markka (FIM) to improve their price competitiveness and to maintain and increase market share. In paper markets, Finnish exporters have pursued a midway pricing strategy aimed at both maintaining market shares and profitability.

Key words:

Forest products, sawnwood, newsprint, export price formation, imperfect competition, elasticity of substitution, law of one price, exchange rate pass-through, Johansen cointegration method, United Kingdom, Germany

Dedicated to my son Tapio

Preface

My research interest on price determination and exchange rate effects on Finnish exports of forest products is related to Finland's participation in the integration of Europe. The realization of European Monetary Union (EMU) on 1.1.1999 is the next step in this process and it will intensely affect the market environment for Finnish forest industry. Although it is difficult to assess the future by analysing historical data, the results of the present study give insights into operating within the impending EMU. The present study is part of the project "Short-term Forecasting System for the Finnish Forest Sector" (MESU) at the Finnish Forest Research Institute (Metla) between 1994 and 1998. The research has been mainly financed by Metla, but research grants from Metsämiesten Säätiö are also gratefully acknowledged.

My special gratitude belongs to Jari Kuuluvainen, who has supervised my work from the beginning with encouragement and guidance. Moreover, as my previous project leader and collaborator, Jari has been an essential support throughout all of my scientific work. I also want to thank Heikki Juslin for his encouragement of my studies and Mikko Tervo for his comments on this summarizing report. My project leader, Lauri Hetemäki, has been important throughout my research and his constructive criticism has greatly increased my skills as a researcher. I am also grateful to Anne Toppinen for her friendship and her participation in this project, as well as her collaboration on other projects. Part of this dissertation is a result from joint research with Anne as well as Pertti Ruuska. Antti Ripatti helped me with methodological questions and the comments of both Ville Ovaskainen and Ashley Selby were also helpful in organizing this summarizing report. I also want to thank, warmly, my pre-examiners, Jussi Uusivuori and Jouko Vilmunen, for their valuable comments and suggestions to improve the summary report. Additionally, Robert Craig has been helpful in improving my English in the final stages of the summary report.

I have been fortunate to share a researcher's life with my beloved husband, Harri, to whom I am profoundly grateful for his support and understanding.

Kerava, October 9, 1998

Riitta Hänninen

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LIST OF SEPARATE STUDIES

The dissertation includes the following separate studies:

Hänninen, R. 1994. Price elasticities of substitution between Finland and the other supplier countries in the demand of the United Kingdom sawnwood imports. *In* Proceedings of the Biennial meeting of the Scandinavian Forest Economics, Helles, F., and Linddal, M. (eds.). *Scandinavian Forest Economics* 35:204-217.

Hänninen, R., Toppinen, A., and Ruuska, P. 1997. Testing arbitrage in newsprint imports to United Kingdom and Germany. *Canadian Journal of Forest Research* 27:1946-1952.

Hänninen, R. H. 1998a. The law of one price in United Kingdom soft sawnwood imports - a cointegration approach. *Forest Science* 44(1):17-23

Hänninen, R. H. 1998b. Exchange rate changes and the Finnish sawnwood demand and price in the UK market. *Silva Fennica* 32(1):61-73.

Hänninen, R. and Toppinen, A. 1998. Long-run price effects of exchange rate changes in Finnish pulp and paper exports. Forthcoming *in Applied Economics*.

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1. INTRODUCTION

1.1. Background and justification

Substantial changes in the export market environment produce challenges for Finnish forest industry. Competition in world markets and Finland's main market area, Europe, is likely to increase in the future due to new producers and low production costs in, e.g., Southeast Asia (Hetemäki 1997). The realization of the third phase of European Economic and Monetary Union (EMU) on 1.1.1999 removes the exchange rate buffer for Finnish forest industry trade with the other EMU countries. Adjustment difficulties to the new market environment may arise, if exchange rate has an important effect on Finnish exports. A common currency will facilitate price comparisons in the EMU area, which may increase competition between producers in different countries. This study produces information about price formation and competition in export markets, which is useful for policy makers and for industry in adjusting to future market changes. For those modeling Finnish forest industry exports, the information is also useful in choosing a suitable modeling strategy.

The study focuses on the two most important export countries for Finland in the European Union, the United Kingdom (UK) and Germany (BDR). They accounted for 32 percent of sawnwood exports, 42 percent of newsprint exports and 60 percent of chemical pulp exports from Finland in 1996 (Statistical Yearbook...1997). Finnish sawnwood and newsprint industries are clearly export industries; the greatest part of their production is exported. In the Finnish chemical pulp industry, the export share of total production is, however, only about 25 percent. Finland's main competitors in paper and pulp markets are Sweden, Canada and the USA, and in sawnwood markets Sweden, Canada and Russia. Sawnwood, newsprint and chemical pulp were chosen for this analysis, because they form relatively homogenous product groups, a prerequisite for an analysis based on unit prices.

In the present study of price formation three different aspects are examined. First, the study examines price substitution. Second, it tests the existence of price differences (the law of one price, LOP) between Finland's and its competitor's products. Third, the magnitude of exchange rate changes on Finnish export demand and prices expressed in terms of foreign

currency (pass-through effect) is examined. The three topics have implications as to the degree of competition in the market. High substitution elasticities, the existence of LOP and a low pass-through (PT) effect would imply perfectly competitive markets, whereas, relatively low substitution elasticities, the rejection of the LOP and a high PT would indicate imperfect competition.

Previous studies on export price formation in the whole of Finnish industry suggest imperfect competition at the aggregate level (Honkatukia 1994). For paper industry exports, the results gave clear support to a judgement of imperfect competition, but for wood products the evidence was less clear-cut (Sukselainen 1986, Honkatukia 1994 and 1995). Because forest industry produces many different products and product groups, it is natural that each of these may show varying degrees of competition in export markets. Thus, it is difficult to obtain a uniform view over the degree of competition in their international trade. Traded products may be relatively homogenous, suggesting more perfect than imperfect competition. However, even in this case market imperfections remain possible.

Studies modeling the trade in different forest products have used different assumptions concerning competition. In some studies the starting point has been perfect competition, while the others assume imperfect competition. The assumptions are not, however, usually explicitly tested. In particular, the differences between Finland's and its competitor's prices have not been statistically tested in earlier works, although there do exist a few studies on the effect of exchange rate on Finnish forest product exports (e.g. Uusivuori and Buongiorno 1990, Vesala 1992, Laaksonen 1998). In the estimation of exchange rate effects, the present study uses Johansen's method (1988), which takes account of time series properties and cointegration of the data. This has been done only in the most recent studies of the Finnish forest sector (e.g., Ripatti 1990, Toppinen and Kuuluvainen 1997, Laaksonen 1998, and Toppinen 1998). Using new estimation methods, additional information is generated on export price formation, thus increasing the understanding of the functioning of forest products export markets.

1.2 Aim and outline

The study examines forest products price formation in Finnish export markets by applying partial equilibrium models and the econometric method. The analysis includes five separate studies concerning soft sawnwood, newsprint and chemical pulp exports to Finland's most important export countries, the United Kingdom and Germany. The analysis can be divided into three specific sub-topics:

1. Price substitution in the UK sawnwood imports from Finland, Sweden, Canada and Russia (Hänninen 1994).
2. The existence of long-term price differences in UK sawnwood imports and in British and German newsprint imports from major suppliers (Hänninen 1998a, Hänninen *et al.* 1997).
3. Exchange rate effects on Finnish sawnwood export demand and price in the UK market and on Finnish newsprint and chemical pulp prices in the UK and Germany (Hänninen 1998b, Hänninen and Toppinen 1998).

The study of the first topic aimed to estimate own-price and substitution elasticities between the four supplier countries' sawnwoods in the UK market. For this purpose, an econometric model for UK sawnwood demand from the supplier countries generalizing the original export demand model of Armington (1969) was specified and estimated using annual data from the period 1961-1990. The studies of the second topic examined differences between Finland's, and its competitor's export prices for sawnwood and newsprint in the UK and Germany. The analyses were done by testing the law of one price (LOP) in these markets using quarterly data from the period 1978-1992 for sawnwood and 1980-1994 for newsprint. The studies of the third topic aimed to examine the effects of exchange rate changes on Finnish sawnwood, newsprint and chemical pulp export prices. For newsprint and pulp exchange rate PT was estimated from a markup price equation (e.g., Hung *et al.* 1993) and for sawnwood, the model was enlarged to include an export demand equation assuming constant elasticity of substitution between supplier countries. The estimations were made using quarterly data from the period 1978-1994 for sawnwood and 1980-1994 for newsprint.

The outline of the present summary paper is as follows. Chapter 2 presents the development of Finland's and its competitor's market shares and the prices of their exports in the UK and Germany. Chapter 3 reviews the related literature and presents the theoretical framework of the studies. Data is summarized in Chapter 4 and Johansen's cointegration method and the testable hypothesis of the study are presented in Chapter 5. Short summaries of each study are presented in Chapter 6 followed by conclusions with suggestions for further research in Chapter 7.

2. EXPORTERS' MARKET SHARES AND UNIT PRICES IN THE UK AND GERMANY

The main exporters to the UK sawnwood market, Finland, Sweden, Canada and Russia, account for about 80 percent of total UK sawnwood imports. The UK is highly dependent on imports, even though it doubled domestic sawnwood production between 1980 and 1994. The percentage ratio share of domestic production to imports has increased during the past 14 years from 18 percent in 1980 to 25 percent in 1994 (FAO 1991 and 1994).

In the newsprint markets of the UK and Germany, the main suppliers are Finland, Sweden and Canada. These three countries account for about 85 per cent of newsprint imports into the UK and 68 percent into Germany. Both countries have doubled their newsprint production between 1980 and 1994 and increased the percentage ratio of domestic production to imports. In 1980 the ratio was 34 percent in the UK and 70 percent in Germany. In 1994, the respective ratios were 45 percent and 116 percent. (FAO 1991 and 1994). The UK is far more dependent on imported newsprint than Germany.

The import shares¹ of suppliers show relatively large business fluctuations in UK sawnwood and newsprint imports over time, while in Germany, market shares have been rather stable (Figures 1-2). The market shares of suppliers (see Table 1 in: Hänninen 1998a and Hänninen et al. 1997) have also been relatively similar indicating that the supplier countries have not differed much from each other in terms of their potential market power.

The main events in the UK sawnwood market were the decrease of the Russian share after the collapse of the Soviet Union and large fluctuations in the Canadian and Swedish shares (Figure 1). Finland's share, however, has been rather stable. In UK newsprint imports, Finland has lost somewhat of its share, while Swedish market share has increased (Figure 2). The decline of the Finnish share in the UK has partially been offset by the acquisition of

¹ Suppliers' percentage import shares are calculated by dividing the sawnwood and newsprint import quantity of each supplier by the respective total imports to the UK and Germany. The data for Figures 1 and 2 are obtained from the British and German national customs statistics (CSO, Statistisches Bundesamt).

production capacity by Finnish forest industries in the UK market during the 1980s. In the 1990s, Finnish firms have also established newsprint production capacity in Germany.

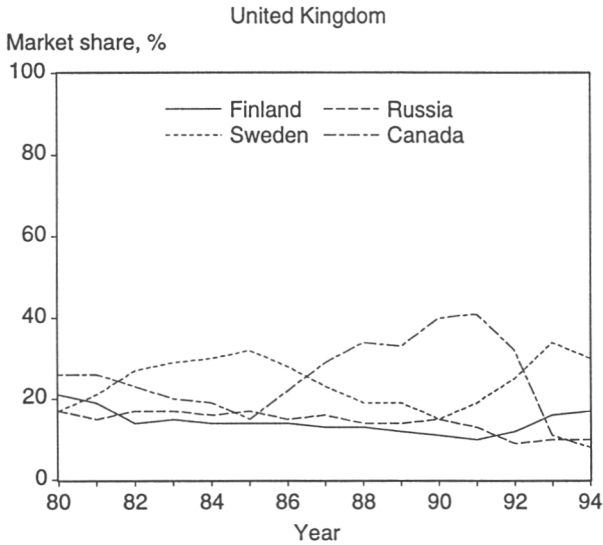


Figure 1. Supplier countries' percentage shares of total sawnwood imports into the United Kingdom, 1980-1994.

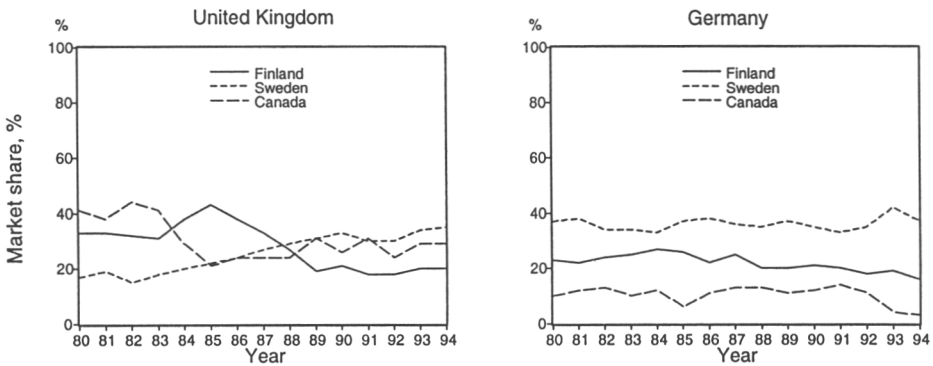


Figure 2. Supplier countries' percentage shares of total newsprint imports into the United Kingdom and Germany, 1980-1994.

Supplier countries' export prices for newsprint and sawnwood are presented in Figures 3-4. If suppliers' prices in the export markets differ, suppliers must have at least some market power.

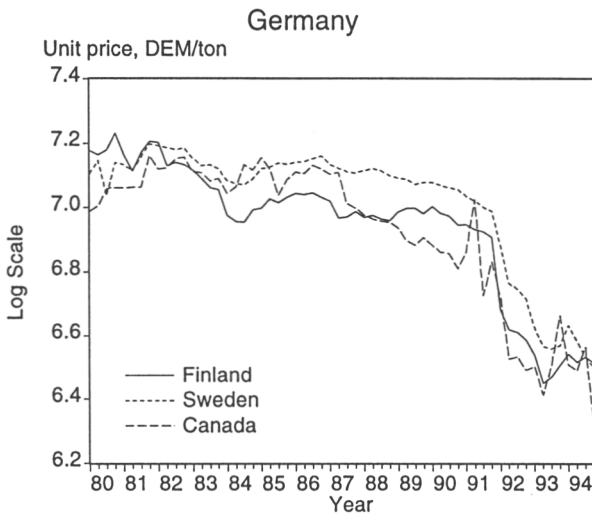
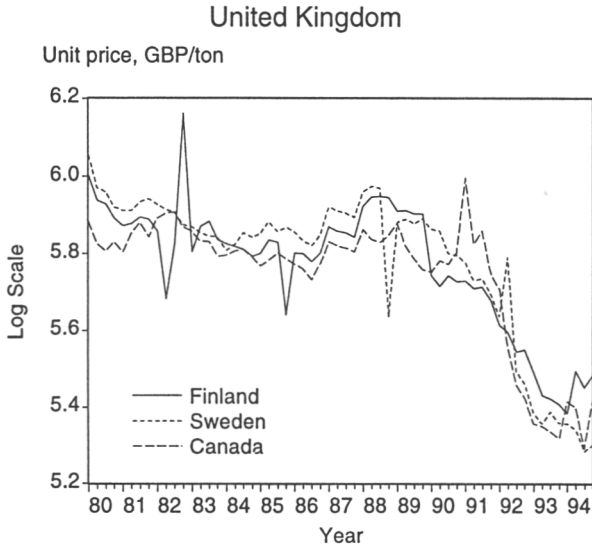


Figure 3. Logarithms of quarterly real unit prices of newsprint imports into the United Kingdom and Germany (GBP/metric ton and DEM/ton), 1980-1994.

From the graphs of the price series², it could be concluded that the suppliers' real unit prices have developed, relatively uniformly, in newsprint markets. After a stable price development during the 1980s, there has been a fall in real newsprint prices in both importing countries during the economic recession of the early 1990s (Figure 3).

Price development in the sawnwood market has been different. Graphs of real sawnwood prices reveal that, especially in the second half of the period, price series have clearly deviated from each other in the UK (Figure 4). Therefore, the LOP is more probably valid for newsprint than for sawnwood. However, explicit tests taking into account time series properties of the data are required to obtain statistically accurate information about the validity of the LOP.

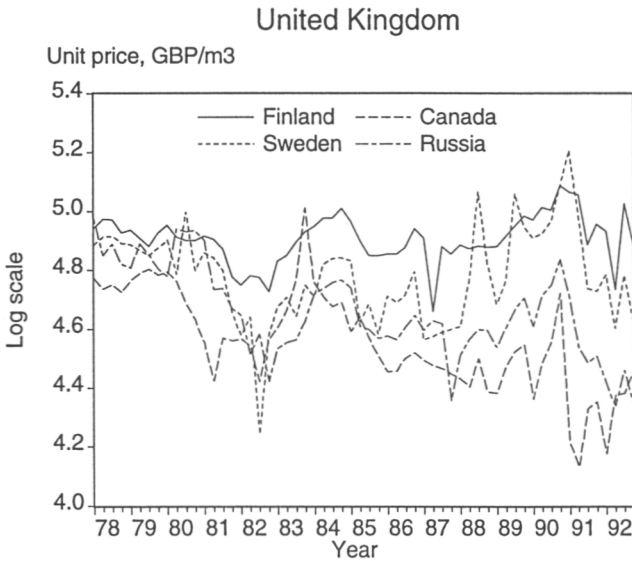


Figure 4. Logarithms of quarterly real unit prices of sawnwood imports to the United Kingdom (GBP/m³), 1978-1992.

² The time series of prices are suppliers' quarterly import unit values into the UK and Germany calculated by dividing the value (CIF, including cost, insurance and freight) of the suppliers' exports by the respective quantity. The sources of the data are British and German national customs statistics (CSO, Statistisches Bundesamt). Unit prices were deflated by the UK and German producer price indexes obtained from the Main Economic Indicators (OECD). More information on the data can be found in Hänninen 1998a and Hänninen et al. 1997.

Graphs of Finnish newsprint prices, and exchange rates with the destination countries (see Hänninen and Toppinen 1998: Figure 1) indicate that these series are negatively correlated with each other. This suggests some degree of exchange rate PT in both markets; FIM depreciations have lowered the price of exports from Finland and FIM appreciations have raised it. This also seems to hold for Finnish sawnwood exports (Hänninen 1998b: Figures 1C and 1D).

During the period studied, the FIM fluctuated widely with respect to GBP and Deutsche mark (DEM) (Hänninen and Toppinen 1998, Figure 1A). Moreover, the exchange rate regime was revised several times. During the period 1978-1991, the FIM was expressed with a currency basket index which was a trade weighted average of the exchange rates of Finland's most important trade partners. In 1984, the peg was changed when the Soviet currency was removed from the index. The currency basket index was supposed to be kept within a band or a target zone. Under this system the devaluations and revaluations were made by specific decision of Finnish government. The FIM was revalued in 1979, 1980, 1989 and devalued in 1978, 1982. In June 1991, the FIM was linked to ECU basket upon application of EU membership. However, as a result of speculative attacks the FIM was devalued in November 1991. During 1992, the FIM with several other European currencies came again under massive speculative attacks and the Bank of Finland decided to abandon the peg to ECU. In September 1992, the FIM was devalued and it was allowed to float until October 1996, when Finland decided to join the European Exchange Rate Mechanism (ERM).

3. THEORETICAL FRAMEWORK AND EARLIER LITERATURE

3.1 Trade models and market competition

The traditional theory of foreign trade explains trade entirely with differences between countries, especially differences in their factor endowments. It assumes comparative advantage, perfectly competitive markets, homogeneity of goods and constant returns-to-scale technology (e.g. Helpman and Krugman 1985). However, traditional trade theory cannot explain why industrial countries have large exports and imports of relatively similar products. In reality, many industries cannot be characterized by constant returns-to-scale. Modern approaches to trade theory explain that economies of scale (i.e., increasing returns-to-scale) in industrial production technologies will lead countries to specialize in production and then trade with one another. It also assumes imperfect competition in the market.

Large forest sector models are commonly based on the assumption of perfect competition³. For example, ETTS V presents forecasts for different countries' forest products exports and imports in this framework (Brooks *et al.* 1995). The Swedish forest sector model assumes perfect competition in West European markets (Byström and Eriksson 1991). The same assumption is also used in partial equilibrium model for the global forest sector, The General Trade Model, GTM, (Dykstra and Kallio 1987) and in the annual medium-term econometric model for the Finnish economy (KESSU) (Hetemäki and Kaski 1992). The quarterly model of Finnish economy (BOF4) assumes perfect competition for Finnish industry exports in the long-run, but imperfect competition in the short run (The BOF4, Quarterly Model...1990). Perfect competition is also assumed in some small partial-equilibrium models of trade in individual forest products. For example, Buongiorno and Gilless (1984), who modelled newsprint trade, Andersen (1993), who studied European newsprint markets, and Brännlund *et al.* (1982), who modeled Swedish exports of forest products to Western Europe, assumed perfect competition.

³ In a perfectly competitive market there are many buyers and sellers and none of them has a large market share. Sellers are price-takers and they can only adjust their quantity at the current world market price. Free entrance to the market and free exit also characterize the existence of perfect competition.

On the other hand, several other studies have assumed imperfect competition⁴. Some previous models of the newsprint industry in North America (e.g., Dagenais 1976, Booth *et al.* 1991) assumed oligopoly with dominant firm price leadership. Brännlund and Löfgren (1995) assumed that Canadian pulp and newsprint industry is a price setter in North America, but a quantity setter in its European exports. In the study of Wiberg (1987), Swedish pulp exports to Germany are assumed to follow price-setting behavior. Oligopoly was assumed by Halonen (1990), who studied Finnish and Swedish pulp and paper exports and Ronnila (1995), who analyzed Finnish paper exports and tested Cournot oligopoly in Europe. Overall, the results indicated the possibility of noncompetitive behavior in the Finnish paper industry. In a number of studies that have applied the theory of Armington (1969), imperfect competition is assumed in the trade of individual forest products. Imperfect competition was assumed, for example, in US softwood lumber imports from Canada (Hseu and Buongiorno 1993), in Finnish sawnwood exports to Europe (Hänninen 1986) and the UK (Tervo *et al.* 1988), in Finnish paper exports to the UK (e.g., Laaksonen *et al.* 1997) and in Finnish paper exports to the UK and Germany (Volk 1983).

In sum, some previous studies of forest product exports have assumed perfect competition, while others have assumed imperfect competition. Forest product markets include many different products and product groups and each of these may show varying degrees of competition in different countries and in different time periods. Thus, it is difficult to obtain a uniform view as to the degree of competition in their international trade. Even though the traded products in question were homogenous, thus suggesting more perfect than imperfect competition, the existence of market imperfections remains a possibility.

In forest industry, there are attributes that would suggest perfect competition. In pulp and paper industry, product standards and specifications are fairly universal, exporters' foreign investment in capacity has increased and mergers and acquisitions are common (e.g.,

⁴ Imperfect market competition is usually characteristic of markets in which there are only a few major producers as well as markets, where each producer's product is seen by consumers as differentiated from those of rivals. In imperfect competition, each producer may be a price-setter, unlike in perfect competition. In imperfect competition, a producer faces a downward sloping demand curve indicating that the quantities sold increase when prices fall.

Canadian Forest Service 1996). In the sawnwood industry, the products of different suppliers are also relatively homogenous.

However, there are also features suggesting market imperfections. The concentration development of forest industry, especially in the paper sector, does not necessarily promote competition. It has also been common to forest industry that firms voluntarily cut production to be able to maintain certain price levels, instead of competing on price (Ronnilla 1995). In the paper industry, the investment costs of new plants are high, which is likely to deter new firms from entering the market. Also, exit from the market may be complicated. Due to the high investment costs, it is not easy to eliminate the excess capacity, if it exists. Exit from the market requires that high-priced production plants are sold to competitors. In sawnwood industry, imperfections in competition can be assumed to arise from regionally concentrated trade and differences in building codes, product dimensions and grades between countries (Canadian Forest Service 1996).

According to Buongiorno *et al.* (1979, p. 643) systematic distortions are also possible in commodity markets in the case where products are technically relatively homogenous. For example, preferences for lumber quality, even of the same species and grade, business practices and habits as well as attachments to a particular supplier country or even to a specific supplier may differ depending on the importing country. These effects are related to switching costs that consumers or other end-users may face when switching between brands or suppliers of a product (e.g., Froot and Klemperer 1989). Okun (1975) stresses the costs of breaking personal sales relationships in industrial transactions. Supplying firms may create relationships with customers by, e.g., using repeat-purchase discounts or making services incompatible with those of other firms.

3.2. Modeling export markets using the Armington approach

The traditional export demand theory of Armington (1969) assumes that products from different countries of origin are imperfect substitutes for each other. The model has been widely applied in studies of international trade of various commodities, for example, US

cotton exports (Babula 1987, Duffy *et al.* 1990), world trade in rice (Ito *et al.* 1990) and fruits and vegetables (Sarris 1983) and for Finland's foreign trade in services (Miikkulainen 1989).

Armington's approach has been applied to the modeling of forest products trade by, e.g., Chou and Buongiorno (1983), who studied the US demand for plywood imports by country of origin. Blatner (1989) estimated the price elasticities for Western European countries' imports of paper products by supplier country. In Finnish studies, the model has been applied in modeling sawnwood exports to western European countries (Hänninen 1986), and to the UK market (Tervo *et al.* 1988) as well as in paper exports (Volk 1983 and Laaksonen *et. al* 1997).

The popularity of the Armington model derives from the simplicity with which it estimates complicated substitutability relationships. The model assumes that the consumers' utility function is homogeneous of degree one in prices and weakly separable so that the consumer's decision process occurs in two stages. First, the total quantity of imports of a product is determined, and then second, the quantity to be imported is allocated among competing suppliers using constant elasticity of substitution (CES) demand functions. Weak separability in consumption (or production) has the advantage that it saves degrees of freedom in the estimation. This makes it possible to reduce the number of variables in the analysis. It is an essential advantage in models of international trade that invariably concern many commodities and several exporting and importing countries. Too many commodities in the consumption function and too many separate inputs in the production function would not lead to a meaningful statistical analysis.

The restrictive assumptions of the model, related to the weak separability, are not usually tested in applications. This has been considered to be problematic in the interpretation of the results (e.g., Davis and Kruse 1993). However, testing is not always possible because of a lack of data. In many studies, when testing is made, the underlying assumptions of the model have not always been satisfied (e.g., Alston *et al.* 1990). This is why some of the applications have refined the original assumptions of the Armington model, for example, by relaxing the homogeneity condition (Ito *et al.* 1990). Also, more general functional forms

and models which account for both non-homogeneity and the existence of variable elasticities of substitution have been presented (e.g. Henning and Martin 1989).

A more flexible functional form, the translog form, presented by Christensen *et al.* (1973), is applied in the present study (Hänninen 1994). This generalizes the original model from two to many factors and relaxes the restrictive assumption of constant elasticity of substitution applied by the Armington model and most of its applications. Thus, it was possible to obtain substitution elasticities for all possible country-pairs separately. Translog cost functions have been common in evaluating factor substitution and economies of scale in forest (and other) industries (e.g., Fuss 1977, Martinello 1985, Puttock and Prescott 1992, Hetemäki 1990).

The present study formulates an export demand model system to estimate substitution elasticities between Finnish, Swedish, Canadian and Russian sawnwoods in the UK market (Hänninen 1994). Because sawnwood is an intermediate product, the model is derived from a cost function of the construction industry and not from the consumer's utility, as in Armington model. The two-stage demand model for sawnwood implies that a representative end-user in the UK first decides how much sawnwood and other (aggregate) inputs are required to produce a certain output. After this, he chooses between sawnwood from different countries of origin, including domestically produced sawnwood.

Assuming cost-minimizing behavior for the sawnwood end user in British construction industry, the optimal sawnwood demands from different origins can be derived from the end user's cost function using Shephard's Lemma. The cost function is homogenous of degree one in prices and it is assumed to be translogarithmic (Christensen *et al.* 1973). Thus, the demand functions in terms of shares can be presented as

$$(3.1) \quad (\partial \ln C / \partial \ln P_i) = P_i X_i / C = S_i = \beta_i + \sum \beta_{ij} \ln P_j,$$

where $S_i = P_i X_i / C$ = the cost share of the i th ($i=1, \dots, 5$) supplier country and i denotes Finland, Sweden, Canada, Russia and the rest of the world. P_j is the respective price of the j th ($j=1, \dots, 5$) supplier. The estimation of the system using the iterative Zellner estimation method (e.g., Pindyck and Rubinfeld 1988) (with the adding up criterion i.e., $\sum S_j = 1$, and the homogeneity and symmetry conditions imposed) provides parameters from which the own-price and substitution elasticities can be calculated for each observation.

The translogarithmic model is more suitable in examining substitution between many countries than the earlier models which assumed constant elasticity of substitution. However, the problems of nonstationary (i.e., trended) data makes the interpretation of the results difficult. As a drawback in using a translogarithmic model, cross-equation restrictions are problematic in the estimation if time series properties are taken into account. For example, in the Johansen method (Johansen 1988) coefficient restrictions between cointegration vectors are not possible.

3.3 Testing the law of one price in forest products markets

The Armington (Armington 1969) model assumes imperfect substitution between products from different countries of origin. This implies that, in the presence of price differences, substitution between supplier countries' products is measurable. The existence of price differences is explicitly studied by testing the law of one price (LOP) in the UK and German markets (Hänninen 1998a and Hänninen *et al.* 1997). The law of one price is a variant of Purchasing Power Parity (PPP). PPP is a long-run relationship and states that, once converted to a common currency, national price levels should be equal. A large amount of empirical evidence on PPP shows that the real exchange rates (nominal exchange rates adjusted for differences in national price levels) tend toward PPP only in the very long run, the speed of adjustment being extremely slow, and that short-run deviations from PPP are large and volatile (Rogoff 1996).

The origins of empirical PPP parity studies trace to the debate on how to restore the world financial system after its collapse during World War I, when gold standard was abandoned

(Rogoff 1996). Prior to the war, the exchange rate between two currencies tied to a gold standard simply reflected their relative gold values. After the war, countries faced the problem of deciding how to reset exchange rates with minimal disruption to prices and government finances. Cassel (1921) was the first to study the subject empirically proposing to use of inflation differentials between countries to calculate the exchange rate changes needed to maintain PPP. Various versions of PPP are used in a vast number of applications, e.g., to choose the right initial exchange rate for a newly independent country, to forecast real exchange rates or to try to adjust price differences in international comparisons of income (Rogoff 1996).

The law of one price is a building block of PPP, and it states that for any good i :

$$(3.2) \quad p_i = er p_c,$$

where p_i is the domestic-currency price of good i , p_c is the foreign currency price and er is the exchange rate defined as the home-currency price of foreign currency. Thus, LOP implies that once prices are converted to a common currency, the same good should sell for the same price in different countries.

The failure of the LOP is found in large number of studies. Of the early studies most notable were Isard (1977), who found deviations from the LOP between U.S., German, Canadian and Japanese exports of goods, such as industrial chemicals, paper and glass. Richardson (1978) found only some evidence of commodity price arbitrage in the exports of industrial products between U.S. and Canada. According to Giovannini (1988) price differentials existed in manufacturing goods. He also found, as Isard (1977) and Richardson (1978), that LOP deviations are highly correlated with exchange rate movements. Also, studies investigating export unit values from a single country to multiple destinations have found deviations from the LOP (e.g., Knetter 1989, 1993).

Several reasons for LOP deviations in traded goods can also be found in the literature. For example, tariffs, nontariff barriers and transportation costs may cause differences between prices in different countries. For some classes of goods, differing national standards may

weaken arbitrage. These different adjustment costs form a buffer within which nominal exchange rates can move without producing an immediate proportional response in relative domestic prices. Deviations from LOP in traded goods may also be explained by "pricing to market" theory (Krugman, 1987 and Dornbush, 1987). In this framework, oligopolistic suppliers can charge different prices for the same good in different countries.

Conventionally, the LOP has been tested by regressing a commodity price in one country on the same commodity price in another country in common currency. Then, the slope coefficient is restricted to equal one and the intercept to equal zero. If the restrictions are not rejected, the law is concluded to hold. The flaw in these tests was the failure to take account of the time series properties of the variables. This was especially criticized in the 1980's and, of the early studies taking the nonstationarity of the variables into account, can be mentioned, e.g., Ardeni (1989), who used a bivariate Engle and Granger cointegration method (1987) to test the LOP.

Single equation methods, however, inefficiently model the interaction between testable price variables and fail to take account of possible dynamics between them (e.g., Edison *et al.* 1997). More recently, systems methods have been developed, e.g., by Johansen (1988). With regard to LOP tests with more than two price variables, single equation methods are unable to find more than one cointegration vector, while according to Goodwin and Grennes (1994), full integration of the markets requires $n-1$ cointegration vectors among n variables.

Although there is a wide literature on testing the law of one price for different commodities, there are few results for forest products. For example, Buongiorno and Uusivuori (1992), tested price differences in US pulp and paper exports to western European countries and Japan and found evidence for the LOP. They used the cointegration approach and a Dickey-Fuller type bivariate method. Johansen's cointegration method was used by Alavapati *et al.* (1997), who examined the price formation of Canadian pulp exports. They found support for the LOP between Canadian and American pulp prices. Jung and Doroodian (1994) used the same method in testing the law of one price between four regional softwood lumber markets in the United States. They interpreted the result as supportive of the LOP, but the

result is problematic, because it does not reveal for which of the four prices the law holds. In the present study, the Jung and Doroodian analysis was extended by conducting additional tests. These tests help to determine to which specific prices the law could possibly apply.

3.4 Exchange rate pass-through on forest product export prices

Pass-through (PT) effects of exchange rates on prices have been analyzed for trade in many commodities and for aggregate imports and exports (see, e.g., the survey by Menon 1995) and the effects are found to vary widely by industry (e.g., Feenstra *et al.* 1996). The variation is related to industry characteristics, such as the degree of competition, product substitutability and the relative (domestic and foreign) market shares (see, e.g., Officer 1986, Menon 1995, Yang 1997).

The large literature on PT arises from the discussion of fixed versus flexible exchange rates in international economics. After the breakdown of the Bretton Woods system of 1971-73, explanations were demanded for the question whether floating exchange rates would play an equilibrating role in trade balances. In this context, the underlying relationship between exchange rates and the prices of internationally traded goods have been widely examined.

Following Dornbusch (1987), two extreme models to measure price relationships in the open economy literature can be presented. One assumes that the law of one price holds in the market (equation 3.1), in which case LOP belongs to PPP literature. The alternative model assumes that products from different origins are imperfect substitutes. Then, the relation between domestic and foreign price levels (real exchange rate) can be presented:

$$(3.3) \quad \phi = p_i / er p_c,$$

where p_i is domestic price level, er is nominal exchange rate and p_c is the foreign price level. Under imperfect competition, price will not equal marginal cost as firms have the possibility to charge a markup on production costs. If the markup is constant, exchange rate

movements change relative prices one-for-one, but, if the markup is variable, the issue becomes how the markup might vary in response to an exchange rate change. The reaction of markups is affected by, for example, the degree of substitutability between domestic and foreign goods and the degree of market integration or separation. The lower the degree of substitutability between goods, and the lower the degree of market integration (the presence of segmented markets), the greater will be the market power of suppliers (Menon 1995).

The degree of the exchange rate effect on prices can be measured by the exchange rate pass-through (PT) coefficient, i.e., the percentage change in export prices (denoted in foreign currency) associated with a one percent change in the exchange rate. The magnitude of the PT may be $0 \leq PT \leq 1$ and it is related to the market structure as follows: If PT is close to zero, competition is perfect in the foreign market and exporting firms have no market power to affect their foreign currency prices as a result of a change in the exchange rate. If PT is unity, pass-through of exchange rate changes to foreign currency export prices is complete and competition is imperfect. PT may be $0 < PT < 1$, indicating incomplete pass-through and imperfect competition. It can be concluded that the closer PT is to unity, the more imperfect is market competition.

However, the interpretation of the relationship between PT and the degree of competition is not always straightforward due to problems in measuring PT. For example, it is possible that PT is zero or near zero in imperfect competition if there are long-term contracts that fix foreign currency prices for months ahead. In this case exchange rate changes can pass through to foreign currency prices only with lag. On the other hand, PT may be higher than zero (incomplete) in perfect competition. For example, producers may increase their export supply as a result of a growth in markups caused by devaluation. The increment of total supply in the export markets usually presses down the foreign price level and the magnitude of PT elasticity becomes higher than zero. This difficulty in interpreting the PT elasticities must be remembered when interpreting the results of the present study.

Unlike other commodities, PT has not been extensively studied for forest products and the few previous studies deal mostly with the USA. Earlier results indicate that PT is higher for wood industry products than for paper products, but conclusions are difficult to draw

because of the scarcity of results. In a study by Vesala (1992), PT for export prices of Finnish paper products was found to be between 0.66 and 0.69 for western Europe, and between 0.16 and 0.30 for the USA. The smaller PT for the USA was assumed to result from the large domestic market in the USA, in which the Finnish share is much smaller than it is in Europe. Alavapati *et al.* (1997) found that the exchange rate effect on Canadian pulp price was small, suggesting that Canadian exporters do not have significant market power in exports. Uusivuori and Buongiorno (1991) estimated the PT for US forest products exports to Europe and Japan. Pass-through was incomplete in most of the product categories. In lumber exports to Japan, PT was high: from 0.79 to 1.04, depending on the species. Menon (1993) also estimated a relatively high PT (0.80) for Australian imports of wood products, but for paper and board, the PT was 0.45.

Many of the previous studies have examined exchange rate effects by estimating the bivariate relationship between the destination currency price of exports and the exchange rate (e.g., Knetter 1989 and 1993, Pick and Park 1991, Uusivuori and Buongiorno 1991). In these studies, as in most other previous studies on PT, estimation is made using Ordinary Least Squares procedures (OLS). Also, OLS in differences has been applied (e.g., Athukolara 1991), and dynamics captured, using polynomial distributed lags (e.g. Hooper and Mann 1989).

In addition to bivariate models, multivariate models have also been applied assuming that the markup is affected by the exchange rate, foreign price and demand pressure. Time series properties of data have usually been taken account in the studies in 1990s (e.g., Athukolara 1991, Athukolara and Menon 1994, 1995, Menon 1993, Hung *et al.* 1993, Naug and Nymoén 1996, Alavapati *et al.* 1997). In earlier studies, based on the markup concept and applying multivariate approaches, single equation models have commonly been estimated (see however, e.g., Kongsted 1996). In contrast, the present study formulates a demand and supply model for estimating PT for sawnwood export price (Hänninen 1998b) that takes into account time series properties and estimates a multiple equation model.

The PT coefficients of the present study were estimated from a markup price equation that is commonly applied in the estimation of PT (e.g., Hung *et al.* 1993). For sawnwood

(Hänninen 1998b) the model also includes an export demand equation derived by assuming a two-stage decision making process, as in the case of Equation 3.1. The functional form assumes constant elasticity of substitution between supplier countries' sawnwoods and the demand and price equations are formulated as

$$(3.4) \quad \text{Export demand:} \quad x_f - x_o = -\eta (p_f - p_o - er) + c + \varepsilon ,$$

$$(3.5) \quad \text{Export price:} \quad p_f = \delta + (1 - \gamma)(er + p_o) + \gamma c_f + u ,$$

where, x_f and p_f are the Finnish quantity and nominal unit price in FIM to the destination market, x_o and p_o are the respective quantity and price (in the importer's currency) of competitors' products, er is the nominal exchange rate (FIM/GPB, FIM/DEM), c and δ are constant terms and c_f is the Finnish unit production cost. ε and u are disturbance terms that describe all the other factors. Parameter η is the constant elasticity of substitution (i.e., the price elasticity of demand) and γ ($0 < \gamma \leq 1$) indicates the PT coefficient ($PT = -(\partial(p_f - er)/\partial er) = \gamma$). The demand equation (3.4) is assumed to be homogenous of degree zero in p_f , p_o and er and the price equation (3.5) homogenous of degree one in er , p_o and c_f . The price relation allows the markup to vary. A constant markup is related to a large PT and a variable markup to a small PT. In the case of a constant markup, exchange rate changes affect foreign currency-denoted prices, while in the case of a variable markup home currency prices are affected. It is assumed that the exchange rate is an exogenous variable in the model.

4. DATA

The main body of the data consists of quantities and unit prices of the supplier countries' sawnwood and newsprint imports into the UK and Germany. The estimation of the substitution elasticities between the main supplier countries' sawnwood in the UK market is based on yearly data from 1961-1990 (Hänninen 1994). The observation period was selected based on the availability of the data. In the rest of the analysis, concerning LOP and PT-elasticities, quarterly (unadjusted) data are used due to the requirements of Johansen's cointegration method (Johansen 1995). The observation period for quarterly analysis was chosen to begin after the years affected by the oil crises in the middle of 1970's so as to avoid the estimation problems caused by large amounts of turbulence in the development of prices. In the following section, the data from all five studies is briefly described.

For the study concerning substitution between supplier countries in British soft sawnwood imports (Hänninen 1994), yearly data on import quantities and values were gathered from Overseas Trade Statistics of the United Kingdom (CSO) using product code SITC 248.2-3. Sawnwood prices for the four suppliers (Finland, Sweden, Canada and Russia) were described by import unit values. They were calculated by dividing the import value (CIF, including cost, insurance and freight) by the respective import quantity, m^3). For the group comprised of the rest of the suppliers of imports the import quantity and unit price were calculated as a residual from the figures of total sawnwood imports.

The studies testing LOP in the UK sawnwood market and in the UK and German newsprint markets also used unit values of imports to describe suppliers' prices (Hänninen 1998a, Hänninen *et al.* 1997). The observation period was 1978-1992 for sawnwood and 1980-1994 for newsprint. The quarterly data on import quantities (m^3 and metric ton) and values (CIF in the destination countries' currencies) were obtained from the national customs statistics: Overseas Trade Statistics of the United Kingdom (CSO) and Aussenhandel Nach Waren und Ländern (Statistisches Bundesamt), where quarterly figures were calculated as sums of the respective months. The product code, for sawnwood, is SITC 248.2-3 and, for newsprint, is SITC 641.1 in the UK data. In the German data, the product code for newsprint is CN 4801 0010 (including item CN 4801 650 for year 1980, item CN 4801 790

for years 1981–87 and item CN 4801 0090 for 1988 onwards). LOP analyses were made by using real price series. The deflators for the nominal unit values were producer price indexes (1985=100) for the UK and Germany obtained from the Main Economic Indicators (OECD).

PT effects were also analysed using unit values to describe prices. For sawnwood, the observation period was 1978-1994. The variables describing sawnwood quantities and prices in British imports were obtained from the same source, using the same product code, as in the study by Hänninen (1998a). However, the data for the years 1991-94 were gathered from EU-trade statistics (European Commission). In order to obtain the Finnish price in FIM (p_f), the Finnish unit price (in GBP) was converted into FIM at the quarterly average nominal exchange rate, er (FIM/GBP), obtained from International Monetary Fund (IMF). The competitors' price (p_o) is the (quantity share weighted) unit price of British imports from Sweden, Canada, and Russia expressed in GBP. The respective quantity (x_o) is the sum of imports from these same three countries. Production costs of Finnish sawnwood, c_f , were described by the production price index (1990=100, manufacture of wood and furniture), obtained from the Bulletin of Statistics (Statistics Finland).

In the study concerning PT for newsprint and pulp (Hänninen and Toppinen 1998), the observation period was 1980-1994. The product codes for newsprint data are the same as mentioned for the LOP study. The codes for pulp (bleached sulphate pulp) data are SITC 251.51 and SITC 251.52 for the UK, and CN 470321 and CN 470329 for Germany. Finnish unit prices of newsprint and pulp (p_f) in FIM were calculated using export values (FOB, free on board) and quantities, which were obtained from the customs statistics of Finland (The Board of Customs). The foreign competitors' prices (p_o) are the average CIF (including cost, insurance and freight) unit values (GBP/ton, DEM/ton) of total newsprint or pulp imports (excluding imports from Finland) into the UK and Germany. The source of competitor data were the customs statistics of the UK (CSO) and Germany (Statistisches Bundesamt). Exchange rates (er) between the Finnish markka and pound sterling and Deutsche mark are quarterly averages obtained from the Bank of Finland. The producer price index (1990=100), for the commodity group (manufacture of paper and paper products) obtained from the Bulletin of Statistics (Statistics Finland) is used to describe

Finnish production costs (c_f) for the pulp and newsprint industry. Because separate indexes are not available for pulp and newsprint, costs of these industries were described with the above mentioned common-cost variable.

The choice of the empirical variables for the study was based on the availability of the data. This means that they do not necessarily describe the theoretical variables of the models well. For example, the prices were described by average unit values because proper data on prices or more disaggregated data on the qualities of each supplier's products were not available. Thus, it was assumed that the suppliers' sawnwood, newsprint and pulp was homogenous enough to allow reliable estimation results. Newsprint and pulp are, in fact, relatively homogenous grades among different paper products. The homogeneity assumption is also justified for Finland, Sweden and Canada because these countries have relatively similar raw material basis, i.e., coniferous wood. For Finnish, Swedish and Russian sawnwood, the assumption is also well justified as producers in these countries have much the same raw material and end-use sectors. Only Canadian sawnwood may not be exactly comparable to the other suppliers' sawnwood.

The use of the production price indexes as a proxy for the domestic cost of the sawnwood, newsprint and pulp industry may cause uncertainty in the testing of PT, but they are used because proper cost variables were not available. Another possibility would have been to describe the production costs by the price of raw material. Quarterly time series for prices of sawlogs and pulpwood are however available only from 1985, which would have made the estimation period too short.

5. COINTEGRATION METHOD

5.1 Johansen's multivariate cointegration method

In the Ordinary Least Squares estimations of many earlier bivariate models, the time series were implicitly assumed to be constant in mean and variance over time or they were assumed to be stationary. However, prices of commodities are usually not stationary, which may lead to a “spurious regression” problem (Granger and Newbold 1974). In this case, the interpretation of the results is difficult, because the conventional t or F statistics calculated for such regressors do not follow standard distributions (e.g., Banerjee *et al.* 1993).

To avoid the problem, time series were often differenced to make them stationary, but this led to another problem; important long-run information of the data was completely ignored. When the idea of cointegration between the variables was developed it became possible to capture long-term effects in dynamic modeling. Granger (1981) showed that, even in the case that all the variables in a vector are not stationary, there may be linear combinations between them that are stationary, i.e., the variables are cointegrated. The cointegration concept is important in economic modeling; if there is an equilibrium relationship between the variables, these variables should not diverge from each other to a great extent in the long run.

In the present study, conventional methods, such as bivariate OLS and multivariate single equation approaches (Engle and Granger 1987) were replaced by Johansen's (1988) cointegration method. It is a convenient way to estimate multiple cointegration vectors and test hypotheses in a simultaneous multivariate framework. This method is further developed by Johansen and Juselius (e.g., 1990 and 1992). Before estimation, the variables of the present study were tested using ADF tests, the results of which indicated that the data for the study are nonstationary (Hänninen 1998a: Table 2, Hänninen *et. al.* 1997: Table 2, Hänninen and Toppinen 1998: Table 1 and Appendix 1 of the present summary paper).

The basic statistical model of Johansen (1988) is the p – dimensional vector autoregressive model with Gaussian errors:

$$(5.1) \ x_t = A_1 x_{t-1} + \dots + A_k x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t = 1, \dots, T,$$

where x_t is a $p \times 1$ column vector of empirical variables, μ is a $p \times 1$ vector of constant terms, D_t is a vector of non-stochastic variables (e.g., seasonal dummies), k is the lag length ($k = 1, \dots, N$) and ε_t is a vector of error terms assumed to be NID(0, Ω). It is assumed that x_t is at most I(1). The purpose of the analysis is to distinguish between stationarity by linear combinations and by differencing. This is why model (5.1) is reformulated in the error-correction form (e.g., Johansen 1995, p.89):

$$(5.2) \ \Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu + \Phi D_t + \varepsilon_t \quad t=1, \dots, T$$

where Δx_t is an I(0) vector. The restriction of μ to cointegration space means that no linear trend is assumed in the data. $\Gamma_1, \dots, \Gamma_{k-1}$ and $\Pi = -I + \Pi_1 + \dots + \Pi_k$ are coefficient matrices. Γ describes short-term dynamics of the process and Π is the matrix of long-run coefficients. The proper lag length, k , for (5.2) can be chosen by diagnostic testing of the residuals of individual equations and by using the Schwarz (SC) and Hannan-Quinn (HQ) information criteria (see Doornik and Hendry 1994). The reformulated model (5.2.) is a traditional first differenced VAR model except for the term Πx_{t-k} .

The matrix Π has a reduced rank in the case of cointegration ($\text{rank } \Pi < p$) and Π can therefore be decomposed into a matrix of loading vectors, α , and a matrix of cointegrating vectors, β , such that $\Pi = \alpha\beta'$. The rank of the matrix Π determines the number of cointegration vectors. When all p components (variables) of x_t are stationary, matrix Π has the full rank p . When no cointegration exists between the variables in x_t , the matrix Π has zero rank. In this case, the variables are integrated of order 1 and the VAR model in differences would be a proper framework for empirical analysis.

The Johansen test generalizes error correction models to a multivariate framework and examines the coefficient matrix Π of the lagged levels to extract information about long-run relationships. Johansen (1988) formulates two likelihood ratio tests, the trace test and the

maximum eigenvalue test, for the rank of Π , which gives the number of cointegration vectors in the set of p variables. The null hypothesis is:

$$(5.3) H_0: \text{rank}(\Pi) \leq r \text{ or } \Pi = \alpha\beta'$$

where α and β are $p \times r$ matrices. The likelihood ratio test statistic for the hypothesis that there are at most r cointegration vectors is:

$$\text{Trace}(r) = -T \sum_{i=r+1}^p \ln(1 - \lambda'_i)$$

where λ'_i s are the squared canonical correlations, i.e., eigenvectors ($\lambda'_1 > \lambda'_2 > \dots > \lambda'_p$) between the two sets of residual vectors R_{0t} and R_{1t} . Residual R_{0t} is obtained from regressing (Δx_t) on the lagged differences and the deterministic variables $(\Delta x_{t-1}, \dots, \Delta x_{t-k+1} D_t)$ and residual R_{1t} is obtained from regressing (x_{t-k}) on $(\Delta x_{t-1}, \dots, \Delta x_{t-k+1} D_t)$. The resulting eigenvectors $V' = [v'_1, \dots, v'_p]$ are usually normalized as $V' S_{11} V = I_p$. The estimators of cointegration vectors β are $\beta = [v'_1, \dots, v'_r]$, the eigenvectors corresponding to the r largest eigenvalues. The ML estimator of the cointegration space β reflects the r largest canonical correlations between the stationary and nonstationary parts of the system. It is important to note that the estimation method produces only the cointegration space. Thus, it is possible to give economic interpretation to the cointegration vectors only after identification.

In testing the rank by the trace test, the null hypothesis is tested in each stage, against the general alternative that there are at most r cointegration vectors. The maximum eigenvalue test is given by the maximal likelihood statistics. The null hypothesis, that there are r cointegration vectors, is tested against the alternative that the rank is $r+1$: $H_0: \text{rank}(\Pi) = r+1$. The test statistic is $L_{\max}(r) = -T \ln(1 - \lambda'_{r+1})$.

The asymptotic distributions of these likelihood ratio tests can be presented as multivariate versions of the Dickey-Fuller distribution and the distributions are also provided by Johansen and Juselius (1990). The present study applies the critical values at five percent

significance level obtained in the connection of the estimation by the FIML econometric program (Doornik and Hendry 1994).

When the number of cointegration vectors, r , has been determined, it is possible to test hypotheses under r by imposing linear restrictions on the matrix of cointegration vectors, β ($=(\beta_1, \dots, \beta_r)$) and loadings, α (Johansen and Juselius (1990, 1992)). α_{ij} measures the weight with which the j th cointegrating vector (β_j) enters the i th equation of the system. Hypothesis testing is done by estimating the model (Equation 5.2) with restrictions and comparing the results to those of the unrestricted model. The tests for these linear restrictions are asymptotically χ^2 distributed. The critical values used in this study were obtained in the estimation by the FIML econometric program at 5 percent significance level (Doornik and Hendry 1994). Under r , the Johansen method (Johansen and Juselius 1990, 1992) formulates different types of restrictions on β or α :

$$(5.4) \beta = H_1\varphi, \text{ where } H_1(p \times s), \varphi(s \times r) \text{ and } r \leq s \leq p$$

$$(5.5) \beta = (H_2, \psi), \text{ where } H_2(p \times r_1), \psi(p \times r_2) \text{ and } r = r_1 + r_2$$

$$(5.6) \alpha = A\psi, \text{ where } A(p \times m), \psi(m \times r) \text{ and } r \leq m \leq p.$$

The alternative hypothesis for (5.4–5.6) is (5.3). Symbol p is the dimension of the model and r is the number of cointegration vectors. Matrices $H_1(p \times s)$, $H_2(p \times s)$ and $A(p \times m)$ are the design matrices, where the restrictions on the parameters $\beta(p \times r)$ and $\alpha(p \times r)$ are formulated by the researcher. The restrictions reduce the parameters to $\varphi(s \times r)$ and $\psi(m \times r)$, where $r \leq s \leq p$ and $r \leq m \leq p$.

In the hypothesis (5.4) the $(p-s)$ restrictions are placed on r cointegration vectors. The test statistic is asymptotically χ^2 distributed with $r(p-s)$ degrees of freedom (Johansen 1988). The hypothesis (5.4) is applied here to test the exclusion of a variable from the cointegration relation and price homogeneity. Hypothesis (5.5) is used to test the stationarity of variables. In this test the null hypothesis is stationarity, unlike in the Dickey-Fuller (1979) test. In (5.5), r_1 of the r cointegration vectors are known *a priori*. Restrictions for these known vectors are presented in the design matrix H_2 and the remaining r_2

cointegration vectors (included in matrix ψ) are unrestricted. The test statistic is asymptotically χ^2 distributed with $(p - r)r_1$ degrees of freedom.

Weak exogeneity of variables with respect to long run parameters α and β can be tested using hypothesis (5.6). Weak exogeneity is required for efficient inference (estimation) of a conditional model. Whether a variable is exogenous depends upon whether it can be taken as “given” without losing information for the purpose at hand. Weak exogeneity of the l th variable for the long-run parameters β implies that all the elements of the l th row of α are zero, i.e., no lagged levels appear in that equation. The test statistic is χ^2 distributed with $r(p-m)$ degrees of freedom.

5.2 Testable hypotheses

In the UK sawnwood market, the LOP was tested between Finnish (pf), Swedish (ps), Canadian (pc) and Russian (pr) prices of exports ($p = 4$) and in the newsprint market between Finnish, Swedish and Canadian prices ($p = 3$). For example, in the case of four price variables, LOP can be written $pf = ps = pc = pr$ in equilibrium. This can also be presented as $pf = ps$, $ps = pc$, $pc = pr$. This implies econometrically that the long-run LOP requires three stationary relationships for sawnwood ($r = 3$) and two for newsprint ($r=2$):

$$(5.7) \quad \begin{array}{l} \text{sawnwood: } pf - ps = I(0), \quad ps - pc = I(0), \quad pc - pr = I(0), \\ \text{newsprint: } pf - ps = I(0), \quad ps - pc = I(0) \end{array}$$

To test (5.7), model (5.2), with three seasonals and a constant term was estimated and Johansen’s hypothesis (5.4.) was used. The second stage of the testing was the pairwise test, in the cases where the rank test supported $r = 1$. Under one cointegration vector, the law cannot hold for four (three) prices simultaneously. Instead, it can only hold for two of the prices To examine for which two prices LOP holds it was tested whether the following restriction on the coefficients of the cointegration vector (β_1) holds under $r = 1$

$$(5.8) \beta_{ii} = -\beta_{jj} \quad (i=1,\dots,4 \text{ for sawnwood and } i = 1,\dots,3 \text{ for newsprint, } i \neq j)$$

(Johansen's hypothesis, 5.4). For example, for sawnwood prices (p_f, p_s, p_c, p_r) six design matrices (H_1) of the form $(1, -1, 0, 0)$, $(1, 0, -1, 0)$, $(1, 0, 0, -1)$, etc. were formulated to test the restrictions under $r = 1$ (Hänninen 1998). The rejection of the restriction (by χ^2 test) would suggest the rejection of the LOP for the price pair in question.

Johansen's method was also used in to test the stationarity of sawnwood price variables (Hänninen 1998a) and the weak exogeneity of newsprint price variables (Hänninen *et al.* 1997). In testing stationarity, Johansen's hypothesis (5.5) was used. The H_2 design matrices to restrict the β_{ii} coefficients were of the form $(1, 0, 0, 0)$, $(0, 1, 0, 0)$, $(0, 0, 1, 0)$ and $(0, 0, 0, 1)$ under $r = 1$. Weak exogeneity (hypothesis 5.6) of the supplier countries' newsprint prices in the UK market were tested so as to have an indication whether any of the three prices can be taken as a driving force behind the others. The test was done by restricting the row of the α_{ii} loadings of the respective variable as $(0, 0, 0)$.

The exchange rate PT estimation for sawnwood (Hänninen 1998b), newsprint and pulp (Hänninen and Toppinen 1998) was based on a markup price equation (3.5). For sawnwood the model also included an export demand equation (3.4). In the case that the data is cointegrated, it is possible to identify cointegration vectors, the coefficients of which describe the long run equilibrium relations (3.4) and (3.5). The starting point in the estimation was again to test cointegration rank using model (5.2) with three seasonals and a constant, the data vectors (x_t) consisting of the following variables:

$$\text{sawnwood:} \quad x_t = [x_f, p_f, er, c_f, p_o, x_o],$$

$$\text{newsprint and pulp:} \quad x_t = [p_f, er, c_f, p_o,],$$

where x_f is Finnish quantity, p_f is Finnish price of exports (FIM), er is exchange rate (FIM/GBP, FIM/DEM), c_f is Finnish unit cost, p_o is competitors' price (in GBP or in DEM) and x_o is competitors' quantity. The results of the rank test showed that $r = 2$ for sawnwood indicating two cointegration vectors in the data with coefficients β_{ij} . Let

$$(5.9) \quad \beta_1: [\beta_{11}(x_f) \quad \beta_{21}(p_f) \quad \beta_{31}(er) \quad \beta_{41}(c_f) \quad \beta_{51}(p_o) \quad \beta_{61}(x_o)]$$

$$\beta_2: [\beta_{12}(x_f) \quad \beta_{22}(p_f) \quad \beta_{32}(er) \quad \beta_{42}(c_f) \quad \beta_{52}(p_o) \quad \beta_{62}(x_o)]$$

denote the two vectors. For newsprint and pulp the result of the rank test was $r = 1$ indicating one cointegration vector that can be described by β_2 of the above system (5.9). Next, β_1 and β_2 were tested by restricting their coefficients to identify the model for sawnwood (relations 3.4 and 3.5), and newsprint and pulp (relation 3.5). For this purpose, Johansen's hypothesis (5.4) was applied. For sawnwood (Hänninen 1998b), the following restrictions were first tested simultaneously for the two cointegration vectors normalized with respect to $\beta_{11}(x_f)$ and $\beta_{22}(p_f)$.

$$(5.10) \quad \beta_1: \quad \beta_{11}=1, \beta_{41}=0, \beta_{31}=-\beta_{21}, \beta_{51}=\beta_{31}, \beta_{61}=-1,$$

$$\beta_2: \quad \beta_{12}=0, \beta_{22}=1, \beta_{62}=0, \beta_{42}=-(\beta_{22}+\beta_{32}), \beta_{52}=\beta_{32}$$

Restrictions $\beta_{41}=0$ in β_1 and $\beta_{12}=0$ and $\beta_{62}=0$ in β_2 mean that c_f is excluded in β_1 and x_f and x_o in β_2 . In the demand relation (3.4) ($p_f - p_o - er$) implies that the coefficients of p_o and er are equal (i.e. $\beta_{51}=\beta_{31}$) and that demand should be homogenous of degree zero in p_f , p_o , and er (i.e. $\beta_{31}=-\beta_{21}$). The linear homogeneity of the price relation (3.5) in the nominal variables p_f , c_f and $(p_o + er)$ implies that the coefficients of p_o and er should be equal ($\beta_{52}=\beta_{32}$) and that the coefficient of c_f should equal the difference between the coefficients of p_f and er ($\beta_{42}=-(\beta_{22}+\beta_{32})$). The model for sawnwood was rejected with these restrictions by χ^2 test at 5 percent significance level and testing was continued by relaxing homogeneity in prices in β_1 . The second stage testable restrictions were

$$(5.11) \quad \beta_1: \quad \beta_{31} \neq -\beta_{21}, \text{ others as in (5.10)}$$

$$\beta_2: \quad \text{restrictions as in (5.10)}$$

Structure (5.11) was not rejected by the χ^2 test and the process was continued by testing whether markup can be restricted to unity in β_2 :

(5.12) β_1 : restrictions as in (5.11)

β_2 : $\beta_{42} = -(\beta_{22} + \beta_{32}) = -1$ others as in (5.10)

In the study of Finnish newsprint and pulp exports only the price relation was estimated (Hänninen and Toppinen 1998) and the restrictions were tested in two stages using the formulations (5.10) and (5.12) for β_2 .

6. RESULTS

6.1 Price elasticities of substitution between Finland and the other supplier countries in the demand of the United Kingdom sawnwood imports

(Hänninen, R. 1994, *Scandinavian Forest Economics* 35:204-217.)

The purpose of the study was to estimate own-price and substitution elasticities between Finnish, Swedish, Canadian and Russian sawnwoods. Armington's export demand model (Armington 1969) commonly applied in earlier studies of trade flows was generalized by assuming variable elasticity of substitution between supplier countries. Thus, a translogarithmic (Christensen *et al.*, 1973) cost share equation system imposing homogeneity and symmetry and using yearly observations between 1961-1990 was estimated. The results indicate that the UK demand for Finnish and Canadian sawnwood has been price elastic. The own-price elasticities were -1.62 for Finland and -1.58 for Canada. The elasticity for Sweden was -0.62 and for Russia about zero. The results for Finland are in line with earlier estimates (e.g., Hänninen 1986, Tervo *et al.* 1988).

The Morishima-elasticities of substitution indicate that the supplier countries' sawnwoods are substitutes in British sawnwood imports. The estimated substitution elasticities between Finnish, Swedish and Canadian sawnwood vary between 1.70 and 2.38 depending on the country pair. The fact that the short run dynamics could not be taken into account in the estimation, however, raises some cause for concern.

6.2 The law of one price in United Kingdom soft sawnwood imports - a cointegration approach

(Hänninen, R.H. 1998a, *Forest Science* 44(1):17-23.)

In this study, the law of one price (LOP) was tested for imports of soft sawnwood into the United Kingdom from Finland, Sweden, Canada and Russia using the concept of cointegration. The data are quarterly import unit values from the supplier countries and cover the period 1978-1992. The study used the multivariate cointegration method of Johansen (1988, Johansen and Juselius 1990 and 1992) instead of the bivariate and single equation methods often used in earlier studies.

The existence of the LOP was initially tested simultaneously for all four price series. Because only one cointegration vector was found in the data, the LOP was also tested separately for pairs of prices under cointegration. The results do not support the LOP. Instead, they indicate the existence of differences over the long run between the suppliers' sawnwood prices. The rejection of the LOP differs from the traditional assumption that commodity prices are perfectly arbitrated, with instantaneous exchange-rate pass-through, at least in the long run. It also differs from the results of Buongiorno and Uusivuori (1992), who tested the law for US paper product exports across destinations. However, the results are in accordance, for example, with Isard (1977), who compared prices of different commodities, including paper products, across exporters (USA and Germany). The results also raise an important issue concerning the assumptions to be made regarding competition in sawnwood markets. In studies of forest products markets it would be useful to statistically test the type of competition before modeling.

6.3 Testing arbitrage in newsprint imports to United Kingdom and Germany

(Hänninen, R., Toppinen, A. and Ruuska, P, 1997. *Canadian Journal of Forest Research* 27:1946-1952.)

The long-run law of one price (LOP) was tested for British and German newsprint imports from Finland, Sweden and Canada by using the theory of cointegration. Johansen's multivariate cointegration method (Johansen 1988, Johansen and Juselius 1990 and 1992) was applied in the empirical analysis. The data consisted of quarterly import unit values from the supplier countries for the period 1980–94. One cointegration vector was found in the UK data, while the prices were not cointegrated in Germany. This indicates that the simultaneous LOP is not valid in either market. In the UK LOP was also tested for price pairs. The results indicated the rejection of the law except for a pair of Swedish and Canadian prices. Differences in the results of the two destination markets may be due to fact that Germany is less dependent on the imported newsprint than the UK. Therefore, the price of domestic newsprint production in Germany, which was not included, may affect the price structure of the German market.

In the UK market, the Swedish and Finnish newsprint prices were found to be weakly exogenous, but this was not true for the Canadian price. This implies that the producers of Scandinavia would not respond if Canadian producers change prices away from Scandinavian prices. But if the Scandinavians change their prices, the Canadians would probably follow. Thus, Scandinavian producers dominate the European newsprint market, while Canadian imports to the area are the residual of the market demand. The LOP between Swedish and Canadian prices and the weak exogeneity of Swedish price in the UK can be interpreted to mean that Canadian price is the adjusting part, and the Swedish price is the determining part, of the relationship. The cointegration tests of the present study provide only weak support to the law of one price, indicating that price differences should be assumed in modeling the UK and German newsprint markets.

6.4 Exchange rate changes and the Finnish sawnwood demand and price in the UK market

(Hänninen, R.H. 1998b. *Silva Fennica* 32(1):61-73.)

The study examines long-run exchange rate pass-through (PT) for the Finnish price of sawnwood in the UK market by estimating an export demand and supply model. The data are quarterly and cover the period 1978-1994. Johansen's (1988, Johansen and Juselius 1990 and 1992) cointegration method, that allows for the estimation and identification of system models, is applied in the estimation.

Two cointegration vectors representing the demand and price equations were found in the data. However, in the export demand relation, homogeneity of degree zero in prices was rejected, indicating that the relation does not follow the assumptions of the economic theory. Thus, the coefficients of the estimated long-run demand equation are difficult to interpret. The PT elasticity obtained from the price equation was close to unity (0.92) and it could be further restricted to equal one. The result is inline with the earlier studies that have estimated relatively large PTs for wood products (e.g., Uusivuori and Buongiorno 1991, Menon 1993). The exchange rate elasticity of export quantity, estimated from the demand equation, was also relatively large (6.56). This result does not agree with earlier studies, which indicate that the trade flows of forest products are not very sensitive to exchange rate variations (Buongiorno *et al.* 1988 and Uusivuori and Buongiorno 1990). The previous results, however, only concern exports to the US market. The results of the present study indicate that exchange rate changes have had an important effect on Finnish sawnwood export prices, measured in foreign currency, and on sawnwood export quantities. A large PT also indicates imperfect competition in the export market.

6.5 Long-run price effects of exchange rate changes in Finnish pulp and paper exports

(Hänninen, R., and Toppinen A. 1998. Forthcoming *in Applied Economics*)

The effects of exchange-rate changes on Finnish prices of newsprint and pulp exports were examined in British and German markets. The pass-through (PT) elasticities were estimated from a markup price equation using Johansen's cointegration method (Johansen 1988, Johansen and Juselius 1990 and 1992) and applying quarterly data from 1980–1994.

The results of the study suggest that devaluations have boosted Finnish exports of paper to the British and German markets. According to the magnitudes of the PT elasticities, pass-through of exchange rate changes to foreign currency export prices has been incomplete. Previous studies have also estimated incomplete PT for paper industry (e.g., Vesala 1992, Uusivuori and Buongiorno 1991, Alavapati *et al.* 1997). According to the results of the present study, the PT elasticity for newsprint was 0.60 in the UK and 0.46 in Germany. Thus, about one half of the changes in the exchange rates has passed through to Finnish foreign currency export prices (GBP and DEM) and about one half to Finnish home-currency export prices (FIM). This indicates that Finnish paper exporters appear to have pursued a midway pricing strategy aimed at maintaining market shares and profitability as the exchange rate changes. In pulp exports, the PT elasticity was small in the UK market (0.07) as compared to in Germany (0.68), implying that Finnish pulp exporters' pricing behaviour has been more clearly destination specific as compared to that of newsprint exporters.

7. CONCLUSIONS

The purpose of the present study was to produce new information about the functioning of forest product export markets. The empirical analysis focused on the price responsiveness of the demand for Finland's and competitor's forest products in the UK and Germany. The examination of price substitution between supplier countries' sawnwood in the UK market was based on the yearly data from 1961 to 1990 (Hänninen 1994). The tests for the law of one price (Hänninen 1998a, Hänninen *et al.* 1997) and the analyses of the exchange rate pass-through (Hänninen 1998b, Hänninen and Toppinen 1998) used quarterly data from 1978 or 1980 to the early 1990's.

The implication, of the results is evidence for the existence of imperfect, rather than perfect, competition between supplier countries in sawnwood and newsprint markets. The results also elaborate the important effects that exchange rate changes have had on Finnish forest product export pricing. The elasticities estimated from the econometric model indicate that substitution between competitors' sawnwoods in British imports were not very high. In perfect competition, substitution elasticities would be infinite. The failure of the LOP assumption suggests price differences between supplier countries in British sawnwood market and in British and German newsprint markets. Exchange rate pass-through (PT) was large in Finnish sawnwood exports and relatively large in newsprint and pulp exports. Only in British market was the PT to pulp export price low. In perfect competition, PT elasticity would be close to, or equal to, zero.

The rejection of the law of one price (LOP) differs from the traditional assumption that commodity prices are perfectly arbitrated. Newsprint and sawnwood are relatively homogenous product groups by product composition, so arbitrage could be expected to hold. In reality, there may be several reasons for the deviations from the LOP. For example, deviations may occur if competition is not perfect, if products are differentiated and if the price series describing the good differ by their commodity compositions in different countries (Officer 1986). In addition, the costs of arbitrage may be high in sawnwood and newsprint markets, with established customer relationships playing an important role.

The large exchange rate pass-through (PT) for Finnish sawnwood export price indicates that the exporters have adjusted their prices measured in foreign currency by almost the full change in the exchange rate. In newsprint exports, currency changes are partly reflected in Finnish export prices in foreign currency and partly in the markup over production costs. Thus, the estimated PT coefficients have implications concerning the pricing strategies of Finnish exporters. Sawnwood exporters have tried to maintain their market shares but not necessarily their markups as the exchange rate changes. Newsprint exporters appear to have pursued a midway pricing strategy aimed at maintaining market shares and profitability as the exchange rate changes. The higher PT estimate for sawnwood than for paper implies that Finnish sawnwood exporters have more efficiently made use of depreciations and devaluations of the FIM to maintain their market shares. This also indicates that the sawnwood industry may have more difficulty than the paper industry in adjusting to competition for market share in the EMU (European Monetary Union) environment.

Under the current Exchange Rate Mechanism (ERM) of the European Monetary System, the possibility of using exchange rate changes to improve Finnish competitiveness and market share has more or less disappeared. Even outside the EMU, the scope for national exchange rate policy would be reduced for Finland (Hetemäki *et al.* 1997). Hence, other means should be found to adjust to future price disturbances caused by shocks in demand. To find means for the adjustment is beyond the scope of this study, but an obvious way to adjust to demand fluctuations is the adjustment of production costs. However, the relatively long-term price contracts that exist in the input markets of Finnish industry make rapid adjustment difficult.

Differentiation of Finnish forest industry products via customer orientation, established customer relationships, product development and by value-added products may be equally important. For example, taking into account the environmental issues that have been found to affect consumer behavior in Finnish export markets (Valtonen *et al.* 1995 and 1997) could form an important advantage in international competition. To find efficient ways to adjust in competition for market share is most important in the Finnish sawnwood industry, where the exchange rate PT was higher than in the paper industry.

Even though the results of the present study increase information on export demand and price formation in Finnish forest product markets, various research questions remain to be explored. The results of the study raise issues concerning the assumption of perfect competition in the modeling of forest products trade flows. Thus, it would be important to explicitly test market competition also in other important export countries (e.g., France, the Netherlands). The scope of this study should also be extended to additional forest products (e.g., plywood, printing and writing papers, paperboards, etc.) and to a more disaggregated product groups in order to have a more complete picture about price formation. It would also be necessary to study the effects of other currencies (e.g., US dollar and Swedish krona) that are important in invoicing forest products. Finally, a short term analysis capturing the dynamics of export demand and price formation in export markets could provide interesting new information compared to the long-term elasticities presented here. Studies on short-term linkages between export markets and the domestic roundwood market could also create useful information for decision makers in the Finnish forest sector.

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APPENDIX 1. Augmented Dickey-Fuller (ADF) unit root tests for the data of the study of Hänninen (1998b) with four lags ($k = 4$) including constant and seasonals.

Variables	Test for levels	Test for first differences
Finnish quantity, x_f	-2.39	-3.02*
Finnish price of exports, p_f	-1.58	-3.68*
Exchange rate, er	-1.83	-3.77*
Finnish unit cost, c_f	-2.68	-3.69*
Competitors' price, p_o	-0.09	-2.80
Competitors' quantity, x_o	-1.42	-2.66

Notes: * denotes the rejection of the null hypothesis of nonstationarity. The critical t -value is -2.90 at the 5% level.

Symbols: x_f = quantity of Finland's sawnwood exports to the UK, 1000 m³, p_f = Finnish unit price of exports in the UK (FIM/m³), er = exchange rate (FIM/GBP), c_f = Finnish unit cost (index), p_o = competitor's price (GBP/m³) and x_o = the competitor's export quantity.

PRICE ELASTICITIES OF SUBSTITUTION
between Finland and the other suppliers in the demand
of the United Kingdom sawnwood imports

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SUMMARY

The purpose of the paper is to estimate substitution and own-price elasticities between Finland and its main competitors in the United Kingdom soft sawnwood market. The model consists of a cost minimizing aggregate demand for sawnwood imports to the United Kingdom and of share equations for each supplier country. The resulting estimates are long term elasticities. The results showed, that the demand for Finnish and Canadian sawnwood was elastic with respect to the price, while demand for Swedish sawnwood was inelastic and the own-price elasticity for Russian sawnwood was near zero. The Morishima elasticities of substitution referred to substitutability between supplier countries' sawnwood in the market. The magnitude of the elasticities indicated that it has been easier to substitute competitors' sawnwood for Finnish sawnwood than vice versa.

Keywords: Sawnwood, imports, two-level demand, Morishima elasticity of substitution, translog function

1. AIM AND APPROACH

The sawnwood industry accounts for more than half the value of timber and wood product exports of Finland. The main part of its production is exported, the traditional customer country being the United Kingdom that accounts for 20 per cent of Finland's sawnwood exports. So, the United Kingdom is still an important market area for Finland's sawnwood, even though its share of total sawnwood exports has declined considerably during the last 30 years. Also Finland's market share and export quantities to the United Kingdom has declined. There are probably several reasons for this declining development. It is assumed here, that the reason is Finland's weakened price competitiveness in the market.

The price competitiveness can be measured by substitution elasticity. It tells, how easily a certain commodity can be replaced by other commodities. Here, it enables us to ascertain, how the mix of sawnwood imported from different supplier countries might respond to a rise in one supplier's price. Then, the magnitude of the elasticity shows how much e.g. a rise in Finland's export price with respect to its competitor's price reduces its export quantity, *ceteris paribus*, in a given export market. The price elasticity estimates of sawnwood import demand can be applied in marketing planning. If information is available on the absolute values of the price elasticities, it is possible, for example, to select a suitable price policy for different customer countries so as to control export earnings.

The principal aim of the paper is to estimate substitution elasticities for sawnwood imports from the main supplier countries, Finland, Sweden, Canada and Russia to the United Kingdom. Also own-price elasticities for imports from these countries are estimated. Because price substitution can be measured only if sawnwood of each supplier country has a distinct price, it is assumed that there are price differences between suppliers in the United Kingdom market. In other words, sawnwood from different countries of origin are assumed to be imperfect substitutes.

Research on substitution elasticities of forest products in foreign trade have been carried out especially in the United States. For example, Chou & Buongiorno (1982) studied substitution between imported and domestic plywood and Chou & Buongiorno (1983) studied substitution in plywood imports to the U.S. between certain Asian countries. Chou & Buongiorno (1984) estimated cross price elasticities between forest products from the U.S. and from the rest of the world in the EEC imports. Further, Hseu & Buongiorno (1993) analysed substitution between species in the demand of U.S. softwood lumber imports from Canada. From Finnish studies we mention e.g. Hänninen (1986) who estimated substitution elasticities between sawnwood from Finland and the rest of the world in West European countries and Tervo et al. (1988), who used the same approach in studying substitution in the United Kingdom sawnwood imports.

In the earlier studies, substitution in foreign trade is most often measured by a cross price elasticity that has been estimated from a model of a constant elasticity of substitution. The present paper is based on the research by Hänninen (1993), that analysed substitution by using a more general functional form that doesn't include this assumption. The data of the study consist of annual time series from the years 1961-1990.

The properties of the time-series are analysed in order to make more reliable conclusions about the elasticity estimates. This is not usually done in the previous studies either. Substitution between Finland's and its competitors' sawnwood is measured by using the Morishima elasticities of substitution. Blackorby & Russel (1989, p. 883) have shown, that Morishima elasticity is a better measure than, for example, Allen partial elasticity of substitution when substitution effects are analysed in many inputs' case.

2. THEORETICAL MODEL FOR SAWNWOOD DEMAND

In order to estimate substitution and own-price elasticities, a model for sawnwood import demand to the United Kingdom is formulated. The model is based on the two-level function approach presented by Fuss (1977). The two-level functions are originally discussed in connection with production theory and applied in the studies of production structures and factor demand. To foreign trade this approach was applied first by Armington (1969), who used a two-level utility function. The two stage model has the advantage that it saves degrees of freedom in the estimation, i.e. it is a way to decrease the number of explanatory variables in the estimable equations.

The two-level demand model of sawnwood implies a two stage optimization process for a representative firm in the end use sector of sawnwood. It is assumed, that a representative producer decides first, how much sawnwood and other aggregate inputs are needed to produce a certain output quantity. After this decision, the producer chooses between sawnwood from different countries of origin. In the present paper, only the latter stage of the model is specified and estimated.

Let us assume that a production function of a representative firm using sawnwood as an input can be presented as follows:

$$(1) Q = f(M, O)$$

where:

- Q = gross output
M = aggregate quantity of sawnwood input
O = all other aggregate inputs, e.g. substitutes of sawnwood, capital, labour etc.

The production function (1) is assumed to be well behaved, i.e. it is assumed that the regularity conditions presented by Diewert (1971, pp. 484-485) hold. Applying the assumption of weak homothetic separability, the production function is written in the form:

$$(2) Q = g(M(M_1, \dots, M_N), O)$$

Where $M = M(M_1, \dots, M_N)$ is the aggregate production function for sawnwood using sawnwood from different countries of origin as arguments. The production function exhibits constant returns to scale with respect to M_1, \dots, M_N , and O. So, it is possible to

regard M as the output of a production process that uses sawnwood from different countries to "produce" the aggregate quantity of sawnwood, M , needed in the production of Q .

Because the production function is assumed to be weakly separable, the marginal rate of substitution between sawnwood from different countries, M_i and M_j , is independent of the quantities of the other inputs O demanded. It is further assumed that the sawnwood aggregate (subfunction M) is homothetic in its components. This means that the sawnwood mix from different supplying countries is independent of the total quantity of sawnwood demanded.

If factor prices and output levels are exogenously given and cost-minimizing behaviour is valid, the Shephard duality theorem states that there exists a cost function dual to every production function. Thus, the production technology implied by (2) can be represented by the following cost function, which is also homothetic and weakly separable:

$$(3) C = h(P_M(P_{M1}, \dots, P_{MN}), P_O, Q),$$

where:

- C = total production cost
- P_M = an aggregator function, i.e. price index of sawnwood from different countries of origin ($j=1, \dots, N$)
- P_O = prices of the other inputs
- Q = gross output.

Under constant returns to scale of the production function and linear homogeneity of the respective cost function, the price index of sawnwood has the property:

$$(4) P_M M = (P_{M1} M_1 + P_{M2} M_2 + \dots + P_{MN} M_N) = \sum P_{Mj} M_j,$$

where M and M_1, \dots, M_N ($j=1, \dots, N$) are related to the sawnwood function $M = M(M_1, \dots, M_N)$ as in (2).

This two-stage optimization process should lead to the same result as straightforward cost minimization, given prices P_{M1}, \dots, P_{MN} , and P_O . For this, the price index P_M must depend only on the prices of the sawnwood inputs (e.g. Solow, 1955-56, p. 105):

$$(5) P_M = P_M(P_{M1}, \dots, P_{MN}) = \text{unit cost } C.$$

Since P_M is the price per unit of sawnwood, it is also the cost per unit to the optimizing agent. It is in fact the minimized cost of total imported sawnwood.

The optimal sawnwood mix can be derived from the above cost function (5). This cost function is assumed to be homogenous of degree one and it is also assumed to be translogarithmic (Christensen et al. 1973). So, it is presented in the following translog form:

$$(6) C = P_M = \beta_0 + \beta_F \ln P_F + \beta_S \ln P_S + \beta_C \ln P_C + \beta_{SU} \ln P_{SU} + \beta_R \ln P_R \\ + 1/2 [\beta_{FF} (\ln P_F)^2 + \beta_{SS} (\ln P_S)^2 + \beta_{CC} (\ln P_C)^2 + \beta_{SUSU} (\ln P_{SU})^2 \\ + \beta_{RR} (\ln P_R)^2] + \beta_{FS} (\ln P_F) (\ln P_S) + \beta_{FC} (\ln P_F) (\ln P_C) + \beta_{FSU} (\ln P_F) (\ln P_{SU}) \\ + \beta_{FR} (\ln P_F) (\ln P_R) + \beta_{SC} (\ln P_S) (\ln P_C) + \beta_{SSU} (\ln P_S) (\ln P_{SU}) + \beta_{SR} (\ln P_S) (\ln P_R) \\ + \beta_{CSU} (\ln P_C) (\ln P_{SU}) + \beta_{CR} (\ln P_C) (\ln P_R) + \beta_{SUR} (\ln P_{SU}) (\ln P_R),$$

where:

P_M = unit cost of aggregate sawnwood imports

P_i = sawnwood prices from the supplier countries, $i=F, S, C, SU, R$,

(F = Finland, S = Sweden, C = Canada, SU = Russia, R = the rest of the world)

Assuming that the cost function (6) is twice differentiable and using Shephard's duality theorem, the derivative of cost with respect to price equals the optimum quantity of a sawnwood from a certain supplier country. Cost minimizing behaviour implies (Fuss 1977, p. 9 and e.g. Greene 1990, pp. 527-528) that the demand functions for the suppliers, in terms of shares in the cost of the sawnwood aggregate, take the form:

$$(7) (\partial \ln C / \partial \ln P_i) (P_i / C) = P_i X_i / C = S_i = \beta_i + \sum \beta_{ij} \ln P_j,$$

Where $S_i = P_i X_i / C$ = the cost share of the i th factor and i = sawnwood imports from Finland, Sweden, Canada, Russia and the rest of the world. The respective cost share equations can be presented as:

$$(8) S_F = \beta_F + \beta_{FF} \ln P_F + \beta_{FS} \ln P_S + \beta_{FC} \ln P_C + \beta_{FSU} \ln P_{SU} + \beta_{FR} \ln P_R \\ S_S = \beta_S + \beta_{SF} \ln P_F + \beta_{SS} \ln P_S + \beta_{SC} \ln P_C + \beta_{SSU} \ln P_{SU} + \beta_{SR} \ln P_R \\ S_C = \beta_C + \beta_{CF} \ln P_F + \beta_{CS} \ln P_S + \beta_{CC} \ln P_C + \beta_{CSU} \ln P_{SU} + \beta_{CR} \ln P_R \\ S_{SU} = \beta_{SU} + \beta_{SUF} \ln P_F + \beta_{SUS} \ln P_S + \beta_{SUC} \ln P_C + \beta_{SUSU} \ln P_{SU} + \beta_{SUR} \ln P_R \\ S_R = \beta_R + \beta_{RF} \ln P_F + \beta_{RS} \ln P_S + \beta_{RC} \ln P_C + \beta_{RSU} \ln P_{SU} + \beta_{RR} \ln P_R$$

where:

S_i = cost shares of the supplier countries, $i=F, S, C, SU, R$,

P_i = sawnwood prices of the supplier countries, $i=F, S, C, SU, R$,

(F = Finland, S = Sweden, C = Canada, SU = Russia, R = the rest of the world)

The above demand system is assumed to satisfy the adding-up criterion $\sum S_i = 1$ and to impose the properties of the neoclassical production theory:

$$(9) \beta_F + \beta_S + \beta_C + \beta_{SU} + \beta_R = 1 \\ \beta_{FF} + \beta_{FS} + \beta_{FC} + \beta_{FSU} + \beta_{FR} = 0 \\ \beta_{SF} + \beta_{SS} + \beta_{SC} + \beta_{SSU} + \beta_{SR} = 0 \\ \beta_{CF} + \beta_{CS} + \beta_{CC} + \beta_{CSU} + \beta_{CR} = 0 \\ \beta_{SUF} + \beta_{SUS} + \beta_{SUC} + \beta_{SUSU} + \beta_{SUR} = 0 \\ \beta_{RF} + \beta_{RS} + \beta_{RC} + \beta_{RSU} + \beta_{RR} = 0 \\ \text{and to satisfy the Slutsky symmetry condition } \beta_{ij} = \beta_{ji}, i \neq j.$$

The estimation the equation system produces parameters for the calculation of own-price and substitution elasticities for sawnwood imports from different supplier countries. The robustness of these estimates depends on the robustness of the estimation results. Thus,

the theoretical restrictions imposed by the underlying theory must be tested. In other words, the necessary and sufficient condition for strict quasi-concavity of the cost function in input prices is that the matrix of the computed elasticities (Hessian matrix) must be negative semidefinite at each observation point. (Berndt 1991, p. 476 and 493). Moreover, the underlying cost function must be monotonically increasing in input prices. This means that the fitted cost shares must all be positive. The cost share equation systems must also satisfy the symmetry restrictions, which are tested by using the Wald test (e.g. Berndt 1991, p. 465).

3. THE OWN-PRICE AND SUBSTITUTION ELASTICITIES OF DEMAND

For the translog cost function, the price elasticities can be presented in the following way. The own price elasticity of input demand can be written:

$$(10) \epsilon_{ii} = \sigma_{ii} S_i$$

and the cross-price elasticity of input demand can be presented in the form:

$$(11) \epsilon_{ij} = \sigma_{ij} S_j$$

(see e.g. Puttock & Prescott 1992, p. 1141). In these formulae:

$$(12) \sigma_{ij} = (\beta_{ij} + S_i S_j) / S_i S_j \text{ and}$$

$$(13) \sigma_{ii} = (\beta_{ii} + S_i(S_i - 1)) / S_i^2$$

where:

β_{ij}, β_{ii} = estimated coefficients of the cost share equations
 S_i, S_j = the fitted cost shares for the supplier countries i and j .

The formulae 12 and 13 are the Allen partial elasticities.

The Morishima elasticity of substitution (MES) is used to measure the substitution effects. It can be defined according to e.g. Chambers (1988, p. 96) as:

$$(14) \text{MES}_{ij} = \epsilon_{ij} - \epsilon_{ji}$$

This elasticity of substitution is asymmetric and the classification of inputs i and j as Morishima substitutes or complements depends critically on which one of the input prices changes. According to 14, inputs i and j are Morishima substitutes (i.e. i is substituted for j) if and only if an increase in p_j causes the input ratio x_i/x_j to increase when p_i is fixed.

If the estimate of the Morishima elasticity of substitution is positive, the two inputs in question can be interpreted as Morishima substitutes or inputs substitutable for each other. If, on the contrary, the elasticity is negative, the inputs are Morishima complements. Because the elasticity estimates for the translog function normally differ

at every observation, they are usually computed at the mean data values (e.g. Greene 1990, p. 528).

4. DATA

The estimation of the model requires time series data on quantities, prices and cost shares of sawnwood imports to the United Kingdom by supplier countries. The quantity of aggregate sawnwood input (M) demanded in the end use sector is described by the total imports of coniferous sawnwood to the United Kingdom. Changes in the quality or species composition that may have occurred during the years 1961-1990 could not be taken into account because of a lack of information.

Sawnwood demanded from different countries of origin are described by the imports of coniferous sawnwood from Finland (F), Sweden (S), Canada (C), Russia (SU) and the rest of the countries (R). The quantities (cbm) and unit prices (£/cbm) of imported sawnwood are obtained from the Overseas Trade Statistics of the United Kingdom. The unit prices are based on the CIF (cost, insurance and freight) values. The series were also checked and compared with other sources (FAO: Yearbook of Forest Products, Finntimber: Statistical Yearbook, TTF: U.K. Yearbook, and Finland's official statistics of foreign trade) and corrected where necessary. The import quantities from the rest of the countries (R) are calculated as the residual $M_R = M - (M_F + M_S + M_C + M_{SU})$. The respective import price P_R is calculated as a weighted (weighted by cost shares) average of the countries included in the residual group.

5. ESTIMATION METHOD AND TIME SERIES PROPERTIES

The study estimates the model by using only the cost share equations (8). In order to specify a stochastic framework for estimation, a random disturbance term is added to each cost share equation. The restriction of linear homogeneity is satisfied by dividing the price variables of the cost share equations by the prices of the residual countries (P_R). Before the estimation, the time series properties of the cost share variables (S_i) and transformed price variables (P_i/P_R) were analysed.

From the viewpoint of the statistical validity of the estimation results, it is important to ensure that the time series are stationary i.e. I(0) processes. A time series is said to be (weakly) stationary if its first two moments, i.e. mean and variance, are constant. Many economic time series do not satisfy the stationarity properties, they are often I(1) processes.

The properties of a stationary series I(0) and a nonstationary series I(1) are quite different. Thus, the degree of integration of the dependent and independent variables must match in the equations. If both are I(0) processes, unbiased parameter estimates

can be obtained and the interpretation of the t-values is standard. However, if both are I(1) processes and the error term is I(0), unbiased estimates can be obtained, but the t-values of the parameter estimates do not have standard interpretations. In this case, the dependent (y_t) and independent variables (z_t) are said to be integrated. If y_t is I(0) and some of the dependent variables are I(1), results are not reliable, unless the subset of variables z_t cointegrates to I(0).

In order to make conclusions concerning the robustness of the estimation results, the properties of the time series were examined by using autocorrelation functions, autoregressive processes of the series and normality and stationarity tests. Summing up the test results, the stationarity of the variables required by the estimation method is not fulfilled in the model, the variables seem to be I(1) series (Hänninen, 1993).

The demand system (8) is first estimated equation by equation using the ordinary least squares method (OLS). Then the symmetry restrictions are imposed, and the systems are estimated by using Zellner's seemingly unrelated estimator (ZEF) (Pindyck & Rubinfeld, 1988, pp.331-334). Because the cost shares sum up to unity at each observation over all equations, only four of the five share equations are linearly independent. The common procedure to handle singular systems is to drop an arbitrary equation and then estimate the remaining share equations by using a method that is invariant to the excluded equation. (e.g. Berndt, 1991, pp. 472-473). However, no information is lost, because the parameters of the excluded equation can be calculated by using the remaining equations and the parameter restrictions (9) (see e.g. Berndt, 1991, pp. 472-473).

6. ESTIMATION RESULTS

The Zellner iterative system estimation method with symmetry restrictions produced the following results for the cost share equations:

(15)

Cost share of Finland:

$$S_F = 0.21 \quad -0.17\ln P_F + 0.10\ln P_S \quad + 0.02\ln P_C \quad - 0.02\ln P_{SU}$$

(34.18)* (-3.10)* (1.42) (0.41) (-0.58)

DW=1.67, R²=.30

Cost share of Sweden:

$$S_S = 0.24 \quad +0.10\ln P_F \quad + 0.02\ln P_S \quad + 0.15\ln P_C \quad - 0.21\ln P_{SU}$$

(20.46)* (1.42) (0.15) (2.20)* (-2.86)*

DW=0.36, R²=.05

Cost share of Canada:

$$S_C = 0.18 \quad + 0.02\ln P_F \quad + 0.15\ln P_S \quad - 0.15\ln P_C \quad + 0.03\ln P_{SU}$$

(20.71)* (0.41) (2.20)* (-2.78)* (0.81)

DW=0.85, R²=.30

Cost share of Russia:

$$S_{SU} = 0.21 \quad - 0.02\ln P_F \quad - 0.21\ln P_S \quad + 0.03\ln P_C \quad + 0.16\ln P_{SU}$$

(28.70)* (-0.58) (-2.86)* (0.81) (2.44)*

DW=0.59, R²=.34

where:

$$\begin{aligned} \ln P_F &= \ln(P_F/P_R) = \text{relative price from} && \text{Finland} \\ \ln P_S &= \ln(P_S/P_R) = \text{" " "} && \text{Sweden} \\ \ln P_C &= \ln(P_C/P_R) = \text{" " "} && \text{Canada} \\ \ln P_{SU} &= \ln(P_{SU}/P_R) = \text{" " "} && \text{Russia} \end{aligned}$$

The t values are presented in parentheses under the coefficients and the statistically significant coefficients (at the 5 per cent level of significance) are marked with *. However, the interpretation of the statistical significance of the coefficients is problematic because the variables are I(1) series. The estimated coefficients are unbiased, if all the variables really are I(1) series and if the error terms are I(0) processes. Here, the residuals may be I(0) processes, because they seem to have tendency to return to the mean and the mean seems to be near zero (Hänninen, 1993). However, in order to draw proper conclusions about the stationarity of the residuals they should be tested. Because the system estimation does not maximize the individual equation R², it is not an appropriate measure of goodness-of-fit in an equation system context. Instead, some other measure is needed to evaluate the results (see e.g. Berndt, 1991, p. 486).

The coefficients for the omitted residual countries, S_R, are calculated by utilizing the estimation results of the above presented equations and the restrictions (9). By using the above symbols, the result is as follows:

(16) Cost share of the residual countries:

$$S_R = 0.16 + 0.08\ln P_F - 0.05\ln P_S - 0.05\ln P_C + 0.03\ln P_{SU} - 0.01P_R.$$

The fitted values of all the estimated cost shares are positive, which indicates that the monotonicity condition is satisfied. The computed Wald test value (7.16) indicates, that the symmetry restrictions are not rejected in the model.

6. ESTIMATES FOR SUBSTITUTION AND OWN-PRICE ELASTICITIES

The long-run own-price and substitution elasticities of sawnwood demand from the supplier countries were computed by using the estimation results 15-16 and the formulae 10-14. The own-price elasticities are presented in Table 1.

Table 1: Long run own-price elasticities for sawnwood from different supplier countries, 1961-1990, 1961-1975 and 1976-1990.

Supplier country	Time periods		
	1961-1990 ϵ_{ii}	1961-1975 ϵ_{ii}	1976-1990 ϵ_{ii}
Finland	-1.62*	-1.54	-1.71
Sweden	-0.69	-0.68	-0.69
Canada	-1.58*	-1.69	-1.46
Russia	0.06	0.00	0.13
Others	-0.91	-0.89	-0.89

*) The underlying parameter estimate (from equations 15) is statistically significant at 5 % level of significance.

The resulting own-price elasticities of sawnwood demand from Finland, Sweden, Canada and the residual countries are of the (right) negative sign (Table 1). For Russia the elasticity is negative only in some years of the period (Hänninen, 1993) and its mean values are positive. Before discussing more about the estimated elasticities, it must be reminded that the interpretation of the statistical significance of the elasticities is problematic, because the variables of the cost share equations were not stationary. Nevertheless, it is assumed here, that the computed own-price and substitution elasticities are statistically significant, if the underlying parameter estimates β_{ii} and β_{ij} (in the estimated cost share equations 15) are statistically significant at the 5 per cent level of significance.

According to the estimation results, demand from Finland and Canada appear to be elastic. This means, for example, that a 1 percent rise in the price of Finnish sawnwood would decrease the demand for it by 1.60 percent. Because of price-elastic demand, Finland and Canada could affect their export earnings by pricing policy in the United Kingdom. They could increase their export quantities and total export revenues by lowering the price. Price rises are, however, problematic. If prices are increased, the market shares are lost to the competitors. When increasing costs cause pressure to the prices, one possibility to minimize the impact is to ensure that price changes are in line with those of the competitors. Because Finland's and the other suppliers' price movements seem to have been rather uniform on average, it seems, that Finland has not, however, been able to much benefit from pricing policy. Therefore, the development of production costs as compared to the competitors' costs and exchange rates have an important role in Finland's sawnwood exports.

For the Soviet Union the own-price elasticity estimate is almost 0 and of the wrong sign and it is not statistically significant. Possible explanations for the zero price-elasticity of Russian sawnwood are different motives for sawnwood exports and different way of trading. Sawnwood demand from Sweden and from the residual countries seem to be price-inelastic in the United Kingdom market. The low price elasticity of Sweden's sawnwood indicates that it could have more possibilities for free pricing than its

competitor, Finland. In price-inelastic market it is possible to raise the product price if production costs increase, without losing market share significantly, *ceteris paribus*.

Because Finland seem not to be able to use effective pricing policy to increase its export quantities and revenues in the United Kingdom market, it could try to push the price elasticity of sawnwood more near to zero. Price-inelastic demand could give more possibilities in free pricing when the production costs are increasing. The price elasticity of the product can be decreased by marketing activities, for example, by different promotional strategies and branding that make the product more unique in the eyes of the customers.

The Morishima elasticities of substitution (MES), obtained here, are not necessarily statistically significant (Table 2). However, some inferences, based on their computed values, are made below. According to the estimates, substitutability is the dominating feature. Previous studies also indicate substitutability between Finland and the other suppliers' sawnwood in the United Kingdom market (e.g. Hänninen 1986; Tervo et al. 1988). For Russia the MES estimate could not be computed, because its own-price elasticity was of the wrong sign (Table 1). The substitution elasticities calculated for different periods (Table 2) indicate that their values have not changed much during the period studied.

Table 2: Long run Morishima elasticities of substitution (MES) for sawnwood from different supplier countries, 1961-1990, 1961-1975 and 1976-1990.

MES between supplier countries	Time periods		
	1961-1990 MES_{ij}	1961-1975 MES_{ij}	1976-1990 MES_{ij}
Finland / Sweden (FS)	1.40	1.36	1.44
" / Canada (FC)	1.85	1.94	1.75
" / others (FR)	1.47	1.44	1.50
Sweden / Finland (SF)	2.22	2.15	2.29
" / Canada (CS)	2.38	2.47	2.29
" / others (SR)	0.85	0.86	0.85
Canada/ Finland (CF)	1.91	1.85	1.97
" / Sweden (CS)	1.70	1.79	1.60
" / others (CR)	0.82	0.79	0.84
Others / Finland (RF)	2.30	2.26	2.34
" / Sweden (RS)	0.61	0.60	0.63
" / Canada (RC)	1.48	1.56	1.41

It can be seen from Table 2 that the Morishima elasticities of substitution are not symmetric. For example, a 1 per cent rise in the price of Finnish sawnwood would increase the relative quantity demanded from Sweden (M_S/M_F) by about 2 per cent. On the other hand, a 1 per cent rise in the price of Swedish sawnwood would increase the relative quantity demanded from Finland (M_F/M_S) by only about 1.4 per cent. This means that a rise in Sweden's price causes a smaller increase in demand from Finland than the increase in demand from Sweden caused by a rise in Finland's price. From this, we can conclude, that it has been easier to substitute Swedish sawnwood for Finnish sawnwood than to substitute Finnish sawnwood for Swedish sawnwood. When comparing the substitution elasticities between Finland and its other competitors, Canada and the residual countries, the same kind of conclusion can be drawn. In general, when prices have increased, it has been easier to substitute sawnwood from competing countries for sawnwood from Finland, than vice versa:

Sweden/Finland = MES-SF= 2.22 > MES-FS= 1.40
Canada/Finland = MES-CF= 1.91 > MES-FC= 1.85
Others/Finland = MES-RF= 2.30 > MES-FR= 1.47.

One should, however, notice that the interpretation of the statistical significance of MES estimates is problematic and that the differences between the elasticities are rather small.

7. CONCLUSION

The two-level demand system and the translogarithmic function were quite satisfactory approaches in the calculation of the substitution elasticities between Finland and its competitor countries in the U.K. sawnwood import demand. The resulting estimates for the Morishima elasticities of substitution indicate that sawnwood from Finland, Sweden and Canada are substitutes in the markets of the United Kingdom. Compared to the previous research the present study produced the following additional information on the substitution. First, the estimated substitution elasticities vary pairwise between supplier countries, because constant elasticity of substitution was not assumed. Second, the magnitudes of the Morishima elasticities indicate that it is easier to substitute sawnwood from competing countries for Finnish sawnwood, than to substitute sawnwood from Finland for competitors' sawnwood when prices are increasing.

Even though the approach of the study was quite satisfactory in the estimation of the elasticities, the interpretation of the statistical significance of the parameters was problematic because of the non-stationarity of the variables. So, in order to get more robust estimates of the parameters, the estimation method should be developed further. The time series properties of the variables should be taken into account in the estimation and also possible dynamic features of sawnwood import demand of the United Kingdom could be included into the model.

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The Law of One Price in United Kingdom Soft Sawnwood Imports—A Cointegration Approach

Riitta H. Hänninen

ABSTRACT. The law of one price states that prices of homogenous commodities, defined in a common currency, are equal throughout the world. It implies cointegration of prices. In the present study, the law was tested for imports of soft sawnwood to the United Kingdom from Finland, Sweden, Canada, and Russia using the concept of cointegration. The data are quarterly and cover the period 1978–1992. The study used the multivariate cointegration method of Johansen instead of the bivariate method commonly used in earlier studies. The existence of the law was tested simultaneously for all four import prices and separately for pairs of prices. The results do not support the “law” but instead indicate the existence of differences over the long run between different suppliers’ sawnwood prices. This suggests that imperfect competition models should be used in explaining and forecasting UK sawnwood imports. The results also raise an important issue concerning testing for competition in forest products markets, which is not usually done in connection with the modeling of markets. *For. Sci.* 44(1):17–23.

Additional Key Words: Forest products, international trade, Johansen’s method, unit root econometrics.

IN THIS STUDY THE LAW OF ONE PRICE is tested for United Kingdom (UK) sawnwood imports from its main supplier countries, Sweden, Finland, Canada, and Russia. The law of one price is a conventional assumption in studies of commodity trade. It states that each good has a single price, defined in a common currency unit, throughout the world (Isard 1977). A number of studies that modeled forest product markets and trade have assumed the law was correct, but several other studies have allowed for price differences in the market. The validity of the assumption has not usually been tested. However, reliable information about the existence of price differences in the market is important when choosing a suitable modeling strategy.

Allowance is made for price differences between countries of destination or countries of origin, for example, in trade models based on the theory of Armington (1969). The same is true of studies modeling trade in plywood (e.g., Chou and Buongiorno 1983), pulp and paper (e.g., Blatner 1989), and sawnwood (e.g., Castillo and Laarman 1984, Mohd Shahwahid 1991, Hseu and Buongiorno 1993, Hänninen 1994). The existence of price differences has been justified, for example, by Buongiorno et al. (1979, p. 643), who state that lumber quality (even for the same species and grade),

business practices, habits, and attachment to a particular supplier country or specific supplier may differ depending on the country of origin.

The other approach, which assumes the law of one price, is used, for example, in the Timber Assessment Market Model (Adams and Haynes 1980) and in the Global Trade Model (Dykstra and Kallio 1987). Brännlund et al. (1982) used it in modeling Sweden’s trade in different forest products, as did Buongiorno and Gilles (1984) in modeling newsprint trade and Boyd and Krutilla (1987) in modeling lumber trade.

There are several studies testing the law of one price for different commodities, and the results do not often support the law (c.f., Isard 1977, Ardeni 1989, Knetter 1993). However, there are only a few studies testing the law for forest products. For example, Buongiorno and Uusivuori (1992), tested price differences in US pulp and paper exports to western European countries and Japan. They used the cointegration approach and the Dickey-Fuller type bivariate method, as did, e.g., Ardeni (1989), and found evidence for the law of one price. Because simultaneity cannot be taken into account in a bivariate testing method, the earlier results may have suffered from simultaneity problems. Nor does the

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multivariate cointegration method of Engle and Granger (1987) account for simultaneity.

By contrast, in the test procedure presented by Johansen (1988) and Johansen and Juselius (1990, 1992), the price equations can be estimated simultaneously. Jung and Doroodian (1994) used Johansen's procedure in testing the law for four regional softwood lumber markets in the United States and found support for the law. However, they did not resolve the question: for which two of the four prices does the law hold. They found one cointegration relationship, which means that the law of one price cannot be valid for all four prices simultaneously. In the present study, Johansen's procedure was also used, but the Jung and Doroodian analysis was extended by conducting additional tests. These tests help to determine for which specific prices the law could possibly hold. Also, the unit root tests for price series are carried out within the same framework.

The present study tested the law of one price in the UK sawnwood markets for the period 1978 to 1992. The law implies that prices of sawnwood from different countries should be equal in the UK market. If this is true, the market can be characterized as a perfectly competitive market, in which sawnwoods from the supplier countries were perfect substitutes for each other. Unlike earlier studies, a complete set of tests was carried out, and the multivariate method with simultaneous maximum likelihood estimation was used to allow for more reliable statistical inference. Contrary to the traditional assumption and to the results of Buongiorno and Uusivuori (1992), the present results do not support the law.

Testing the Law of One Price

In the present study, the law of one price for UK sawnwood imports was tested using the concept of cointegration for nonstationary time series. The study used the multivariate testing method of Johansen (Johansen 1988, Johansen and Juselius 1990, 1992) instead of the bivariate tests usually applied in earlier studies. It also expanded the test for softwood lumber prices (Jung and Doroodian 1994) by running additional tests in Johansen's framework: the stationarity of the price variables was examined and the law of one price was tested for pairs of prices.

The testing procedure consisted of two stages. First, it was tested whether the law holds simultaneously for Finnish, Swedish, Canadian, and Russian prices. For this purpose, the cointegration rank of the price data was determined. The rank defines how many cointegration vectors, r , can be found in the data. Second, the law was tested for pairs of prices under r .

As an example of the testing of the law, let P_{it} and P_{jt} be sawnwood prices of two supplier countries, expressed in the UK currency. Assuming no transportation costs, the law of one price implies that in equilibrium $P_{it} = P_{jt}$. Then, the law can be stated as

$$P_{it} = h(P_{jt})^b \quad (1)$$

Using lower case letters to denote logarithms and including an error term, Equation (1) can be written as

$$p_{it} = a + bp_{jt} + u_{ijt} \quad (2)$$

The law of one price implies that the constant term, a , should not be significantly different from zero, b should not be significantly different from one, and u_{ijt} should be distributed identically and independently. Because price levels are nonstationary, the ordinary least squares method would not produce reliable inferences. Therefore, the cointegration method was used.

Assuming the law holds, Equation (2) can also be written as the difference between the two prices, $u_{ijt} = p_{it} - p_{jt}$. If u_{ijt} is $I(0)$, the prices are cointegrated and the law holds. Due to this proportionality of prices, their coefficients, β_{i1} and β_{j1} , in the estimated cointegration vector (β_1) should be equal but of opposite sign: $\beta_{i1} = -\beta_{j1}$. This property is used in the following.

In the case of four price variables, which in this study are prices of sawnwood from Finland, Sweden, Canada, and Russia, the law of one price can be written $p_{1t} = p_{2t} = p_{3t} = p_{4t}$ in equilibrium. This can also be presented with three relationships that form the first null hypothesis:

$$H_0: p_{1t} = p_{2t}, p_{2t} = p_{3t}, \text{ and } p_{3t} = p_{4t} \quad (3)$$

Hypothesis (3) was tested by examining whether there are three cointegration vectors in the cointegration space. If this is true, the law holds for all four prices simultaneously.

The second stage of the testing procedure was the pairwise test of the law under r . Because only one cointegration vector was found, the law cannot hold for four prices simultaneously. Instead, it can hold for two of the prices. The second null hypothesis is then

$$H_0: p_{it} = p_{jt} \quad (i = 1, \dots, 4, i \neq j) \quad (4)$$

This hypothesis was tested by investigating whether the restriction $\beta_{i1} = -\beta_{j1}$ is valid for any two prices in the cointegration vector β_1 (restriction A4 in Appendix 1). If the hypothesis (4) is rejected, the law of one price is also rejected for that price pair.

Data

The data for the study consisted of prices of soft sawnwood imported to the UK market from Finland, Sweden, Canada, and Russia. The data were quarterly, seasonally unadjusted, and covered the period from 1978 to 1992. The four supplier countries together have accounted for about 80% of UK sawnwood imports during the period studied, and their individual import shares have been quite similar (Table 1). This means that suppliers have not differed much in terms of market power.

The import prices of sawnwood from Finland, Sweden, Canada, and Russia to the UK were described by the average import unit values (£/m³) based on CIF (including cost,

Table 1. Suppliers' percentage shares of sawnwood imports to the United Kingdom.

Supplier country	Average percentage shares of total import quantity		
	1978-1984	1985-1992	1978-1992
(%).....		
Finland	17	12	15
Sweden	25	23	23
Canada	22	31	27
Russia	17	14	15
Other	19	20	20
Σ	100	100	100

insurance, and freight) figures. Data on quantities and values of sawnwood (SITC 248.2-3) imports were taken from the Overseas Trade Statistics of the United Kingdom (CSO) for the period 1978-1990 and from the intra- and extra-EU trade statistics (European Commission) for the period 1990-1992. The exchange rate, £/ECU, needed for the observations over the period 1990-1992 was obtained from International Monetary Fund (IMF) statistics. The unit price series were deflated by the UK producer price index (1985 = 100) obtained from the Main Economic Indicators (OECD). All the analysis of the study was carried out using real price series in logarithms.

Graphs of the prices of sawnwood show sharp fluctuations during the period studied, 1978-1992 (Appendix 2). Fluctuations in the suppliers' prices have widened since the 1960s, which is presumably due to the two main events in the international financial markets: the breakdown of the Bretton Woods system in 1971-1973 and the oil crises in the middle of the 1970s. After the breakdown of 1971-1973, the international currency system changed; many countries let their currency float, while Finland and Sweden each used a currency index to define the average exchange rate.

Average import unit values were used to describe prices because more disaggregated data on the quality composition of each supplier's sawnwood were not available for the period studied. This means that sawnwood imported from the four supplier countries to the UK is assumed to be homogeneous enough to allow for reliable testing of the law of one price using unit values. For Finnish, Swedish, and Russian sawnwood, the assumption is well justified as producers in these countries have much the same raw material and end-use

sectors. Canadian sawnwood competes with the other suppliers' sawnwood mainly in the market for structural sawnwood. Moreover, tree species of Canadian sawnwood differ from the competitors' species (NUTEK 1992). This means that sawnwood from Canada may not be exactly comparable to the other suppliers' sawnwood, which must be taken into account in drawing conclusions.

Method and Empirical Results

Cointegration and Stationarity of Price Variables

The cointegration method presupposes that the series to be tested are nonstationary unit root processes. The existence of a unit root was first tested using the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979). After the cointegration estimation, the variables were also tested by Johansen's method, and the results were compared. The ADF test results presented in Table 2 are based on the equation including constant, trend, seasonals, and three lags. According to the results, the levels of the price series seem to be nonstationary, while the first differences are stationary. Thus, it is concluded that all the levels of the prices are nonstationary $I(1)$ processes. The inferences regarding nonstationarity are invariant to the different numbers of lags (1 to 5) included in the test equation.

Hypothesis testing was initiated by determining the cointegration rank, r , which defines the number of cointegration vectors in the price data. For this purpose, Johansen's method was applied. Johansen's method uses a statistical model that is a p -dimensional VAR(k) process. It can be reparameterized in error correction form as

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t = 1, \dots, T \quad (5)$$

where Δx_t is a $I(0)$ vector, μ is a vector of constant terms, D_t is a seasonal dummy, and k is the lag length ($k = 1, \dots, N$). The constant term can be restricted to the cointegration space in the estimation if there is no linear trend in the data. $\Gamma_1, \dots, \Gamma_{k-1}$ and $\Pi = -I + \Pi_1 + \dots + \Pi_k$ are coefficient matrices. Π is the matrix of long-run coefficients, and it can be decomposed into a matrix of loadings, α , and a matrix of cointegrating vectors, β , i.e., $\Pi = \alpha\beta'$. The

Table 2. The Augmented Dickey-Fuller (ADF) and Johansen unit root test results for prices of sawnwood with three lags.

Price	ADF ^a test statistic	Decision	Johansen ^b	
			LR test statistic	Decision
Levels				
Finland, <i>pf</i>	-3.14	I(1)	22.25*	I(1)
Sweden, <i>ps</i>	-2.35	I(1)	23.01*	I(1)
Canada, <i>pc</i>	-3.03	I(1)	27.39*	I(1)
Russia, <i>pr</i>	-3.08	I(1)	24.66*	I(1)
First-differenced				
Δpf	-4.04*	I(0)	—	—
Δps	-4.47*	I(0)	—	—
Δpc	-4.07*	I(0)	—	—
Δpr	-3.78*	I(0)	—	—

^a The null hypothesis is nonstationarity. Critical t -value -3.49 at 5% level.

^b The null hypothesis is stationarity. Critical χ^2 value 7.82 at 5% level.

Table 3. Diagnostic tests for residuals of price equations obtained from cointegration estimation of Model (5).

Tests/ standard errors	Finland Δpf	Sweden Δps	Canada Δpc	Russia Δpr
Autocorrelation ^a	7.58	6.04	6.03	8.97
Normality ^b	4.82	3.96	6.21*	0.42
Standard errors	0.07	0.12	0.10	0.07

^a Critical χ^2 value 9.49 at 5% level.

^b Critical χ^2 value 5.99 at 5% level.

cointegration vectors define the stationary linear combinations (long-run relations) of the variables in x_t . ε_t is a vector of error terms assumed to be $NID(0, \Omega)$, i.e., the error terms are normally and independently distributed with expectation zero and variance matrix Ω .

Different values found for r in the cointegration estimation have different implications: (a) If $r = 0$, the price series are not cointegrated and OLS in differences should be used in testing (Banerjee et al. 1993, p. 256); b) if $r = 4$, the series are individually stationary and OLS in levels can be used in testing; (c) if $0 < r < 4$ the price series are cointegrated.

The first hypothesis (3) was tested by determining the number of cointegration vectors, r , in the data. The Model (5) was fitted to the data for the period 1978 through 1992. The dimension of the estimable model is four ($p = 4$). It consists of equations for the prices of sawnwood from Finland, Sweden, Canada, and Russia. Because the restriction of the constant term to the cointegration space did not affect the test inferences, the test results are presented without this restriction.

The diagnostic tests for the cointegration estimation of Model (5) are satisfactory (Table 3). Autocorrelation of the residuals was examined using the Lagrange-Multiplier (LM) test, which is valid for systems with lagged dependent variables. The null hypothesis (no autocorrelation) was accepted with three lags ($k = 3$) at the 5% level; hence three lags were used in the model. Normality of the residuals was tested with the Jarque-Bera test (Jarque and Bera 1980) at the 5% level. The null hypothesis (normality) is rejected for Canada. However, the test value in that case is very close to the critical value, which indicates that nonnormality is not a serious problem. The standard errors of the equations for Sweden and Canada are high, indicating that the model does not fit the price data very well. The high standard errors may be due to sharp price fluctuations (Appendix 2), which are not necessarily easy to model.

The results of the cointegration estimation of Model (5) indicate that $r = 1$. Johansen's method (Johansen 1988) formulates two likelihood ratio tests for the cointegration rank, r , the trace test and the maximal eigenvalue test. Since

Table 4. Test results for number of cointegration vectors (r) in price data for sawnwood.

Null hypothesis	Eigenvalues λ_i	LR-test statistic for trace test	Critical values λ_{trace}^a
$H_0: r = 0$	0.43	53.52*	47.20
$H_0: r \leq 1$	0.25	20.43	29.90
$H_0: r \leq 2$	0.15	9.10	15.40
$H_0: r \leq 3$	0.03	1.63	3.80

^a Critical value at 5% level (Doornik and Hendry 1994).

both tests gave similar results, only the trace test results are presented here. According to the λ_{trace} likelihood ratio test (Appendix 1), the hypothesis of no cointegration ($r = 0$) is rejected, and one cointegration vector is accepted at the 5% level (Table 4). This result implies that Hypothesis (3) of the simultaneous law of one price can be rejected.

The normalized eigenvectors (β_i) and their weights (α_i) obtained from Model (5) are shown in Table 5. Of these eigenvectors, the first relation, β_1 , which is presented in the first column of Table 5, is most highly correlated with the stationary part of the process Δx_t when corrected for the lagged values of the differences. Thus, β_1 is the cointegration vector found in the present price data.

The normalized coefficients, α_1 , in the first column of the weights in Table 5, are the weights with which the error correction term enters the four equations of the system. The α 's are adjustment coefficients that describe the average speed of adjustment toward the estimated equilibrium state. A low coefficient indicates slow adjustment, while a high coefficient indicates rapid adjustment.

After determining the cointegration rank, r , the unit root test was applied for each price variable, under r (r was found to be equal to 1). The null hypothesis was that of stationarity: $H_0: p_{it} = I(0)$, ($i = 1, \dots, 4$). The test was done by restricting the coefficient, β_{i1} , of the testable price, p_{it} , to unity and the coefficients of the other three prices to zero in the cointegration vector β_1 (restriction A3 in Appendix 1). According to Johansen's test, the null hypothesis of stationarity is rejected for all variables (Table 2). Thus, the results of the ADF and Johansen's test are in line with each other and indicate the nonstationarity of the variables. The basic difference in these tests is in respect to the null hypothesis. The ADF test assumes nonstationarity as the null, while Johansen's test assumes stationarity.

Hypothesis (3) is rejected in the present study, because only one cointegration relation was found in the data. This means that the law of one price cannot be valid for all four prices simultaneously. The result $r = 1$ indicates, however,

Table 5. Results from cointegration estimation of Model (5). Normalized eigenvectors (β_1, \dots, β_4) for Finnish, Swedish, Canadian, and Russian sawnwood prices and respective weights ($\alpha_1, \dots, \alpha_4$).

Price variables by country	Eigenvectors				Weights			
	β_1	β_2	β_3	β_4	α_1	α_2	α_3	α_4
Finland	1.00	1.17	1.31	-2.08	-0.55	-0.07	0.04	0.04
Sweden	-0.77	1.00	-0.60	0.08	0.76	-0.13	0.18	0.05
Canada	-0.37	0.96	1.00	0.22	-0.23	-0.14	0.02	-0.06
Russia	0.58	-0.58	-1.04	1.00	-0.44	-0.00	0.15	-0.01

that a stationary relationship between any two price variables is possible. This is why testing of the law was continued by examining pairs of prices.

Testing the Law of One Price for Pairs of Prices

Next, the cointegration vector, β_1 (Table 5), estimated from Model (5) was used to test whether the law of one price holds in pairwise comparisons of prices (Hypothesis 4). The cointegration vector can be interpreted in terms of an error correction mechanism measuring the changes in prices of sawnwood from Finland. The long-term equilibrium relation, β_1 , between price variables can be presented as

$$pf = + 0.77ps + 0.37pc - 0.58pr \quad (6)$$

where *pf*, *ps*, *pc*, and *pr* are prices of sawnwood from Finland, Sweden, Canada, and Russia. In Equation (6), the coefficient of the Canadian price is small compared to the coefficients of the Finnish and Swedish prices. Moreover, the coefficient of Russian price is negative. One reason for the small coefficient of the Canadian price may be in the quality composition of Canadian sawnwood. As was mentioned earlier, sawnwood from Canada may not be quite comparable to the other suppliers' sawnwood in the UK market. The negative coefficient of the Russian price is more difficult to explain. It may be due to the export pricing policy, which has been different from that of the other countries.

The validity of the law of one price was tested under $r = 1$ for two price variables at a time (Hypothesis 4) by assuming that they are proportional. According to Equation (6), magnitudes of the price coefficients differ so widely from each other that proportionality between any two prices is unlikely. Proportionality would imply that the coefficients in the estimated cointegration vector, β_1 (Table 5), would be about the same magnitude and of opposite sign. To test the law by price pairs, Johansen's restriction (A4), presented in Appendix 1, was applied. The Hypothesis (4) was tested by restricting two price coefficients of the cointegration vector at a time to $\beta_{j1} = -\beta_{i1}$. The test results, presented in Table 6, indicate rejection of the null hypothesis for all pairs of prices. Consequently, the law of one price is rejected for the UK sawnwood imports for the four supplier countries.

Table 6. Test results for law of one price by price pairs obtained from Model (5) with restrictions on β_1 under $r = 1$.

Testable price pair	$H_0: P_{ii} = P_{jj}$ ($\beta_{i1} = -\beta_{j1}$)	Test (under $r = 1$) LR test statistic ^a
<i>pf</i> & <i>ps</i>	$\beta_{11} = -\beta_{21}$	21.75*
<i>pf</i> & <i>pc</i>	$\beta_{11} = -\beta_{31}$	29.90*
<i>pf</i> & <i>pr</i>	$\beta_{11} = -\beta_{41}$	29.92*
<i>ps</i> & <i>pc</i>	$\beta_{21} = -\beta_{31}$	28.43*
<i>ps</i> & <i>pr</i>	$\beta_{21} = -\beta_{41}$	27.67*
<i>pc</i> & <i>pr</i>	$\beta_{31} = -\beta_{41}$	24.46*

Symbols: *pf*, *ps*, *pc*, and *pr* are prices of sawnwood from Finland, Sweden, Canada, and Russia and β_{j1} is coefficient of respective price in cointegration vector β_1 .

^a Critical χ^2 value 7.82 at 5% level.

Summary and Conclusion

In this paper the law of one price was tested for the United Kingdom's soft sawnwood imports from Finland, Sweden, Canada, and Russia. Johansen's multivariate cointegration approach was used in the testing. The test was applied using quarterly import unit values deflated by producer prices for the period 1978 through 1992.

The results do not support the law of one price for UK sawnwood imports. Thus, the price differences among supplier pairs did not tend to disappear over the long run. This result differs from the traditional assumption that commodity prices are perfectly arbitrated, with instantaneous exchange-rate pass-through, at least in the long run. It also differs from the results of Buongiorno and Uusivuori (1992), who tested the law for US paper product exports across destinations.

The result do agree, for example, with Isard (1977), who compared prices of different commodities, including paper products, across exporters (US and Germany); Ardeni (1989), who tested the law for certain exporters of agricultural and metal products; and Knetter (1993), who found price differences across exporters for certain agricultural products.

The reasons for the price differences between suppliers may be connected with, for example, exchange-rate pass-through and institutional factors that influence the prices of each country. Moreover, the costs of arbitrage may have been high in sawnwood markets, in which established customer relationships play an important role. However, the price differences between Canada and the other countries may also be a result of the difference in quality composition between the Canadian and other sawnwoods. Canadian sawnwood may not be quite comparable with the other suppliers' sawnwood.

The results of the study suggest that models distinguishing products by country of origin and allowing for price differences between exporters can be applied in modeling UK sawnwood imports. They also raise an important issue concerning the assumptions to be made regarding competition in sawnwood markets. Certain features of western European sawnwood trade support the use of imperfect competitive models. The market can be characterized as having four dominant exporters: Finland, Sweden, Canada, and Russia. This intuitive feature should, however, be tested empirically if it is to be used in modeling markets.

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Appendix 1. Johansen's Testing Method

In the following, Johansen's testing method is briefly summarized (for details, see Johansen 1988, Johansen and Juselius 1990, 1992). The maximum likelihood estimators of the α and β parameters for Model (5) are obtained by solving an eigenvalue problem (Johansen 1988). To determine the cointegration rank (i.e., number of cointegration vectors), Johansen formulates two likelihood ratio tests: the trace test and the maximum eigenvalue test. The null hypothesis can be presented as

$$H_0 : \text{rank}(\Pi) \leq r \quad (A1)$$

Because $p = 4$ in the model, the determination of cointegration rank involves four different stages in testing H_0 . The null hypothesis applied for the trace test is that the number of cointegrating vectors is less than or equal to r , where r is 0, 1, 2, or 3. In each stage the null hypothesis is tested against the general alternative: $r = 0$ against $r \geq 1$, $r \geq 1$ against $r \geq 2$, etc. The hypothesis (A1) is tested using the likelihood ratio test of the form

$$-2 \ln(Q) = -T \sum_{i=r+1}^p \ln(1 - \lambda_i) \quad (A2)$$

where T is the number of observations and the λ_i 's are the smallest squared canonical correlations (eigenvalues). The testing strategy is a multivariate analogue of the Dickey-Fuller (1979) test.

When the number of cointegration vectors, r , is determined by the data, it is possible to test different hypotheses by restricting the cointegration vectors, β , or their weights, α . Testing is done by estimating Model (5) with restrictions and comparing the results to the Model (5) without restrictions. The test statistics are asymptotically χ^2 distributed.

Under r , the Johansen method (Johansen and Juselius 1990, 1992) formulates the following restriction on β , which can be used to test the stationarity of a variable:

$$\beta = (H, \psi) \quad (A3)$$

where $H(p \times r_1)$, $\psi(p \times r_2)$ and $r = r_1 + r_2$

In (A3), p is the dimension of the model (here $p = 4$), r is the number of cointegration vectors, of which r_1 vectors

are known *a priori*. Matrix H is a design matrix, where the restriction on the β matrix is formulated by the researcher. The test can be done by estimating Model (5) with restrictions, so that r_1 cointegration vectors are restricted, and the remaining r_2 cointegration vectors (included in matrix ψ) are unrestricted.

Because the number of cointegration relations was found to be one, $r_2 = 0$ and $r_1 = 1$ in the present study. Stationarity is tested by restricting the coefficient of the testable variable (β_{i1}) to unity and the other coefficients to zero in the cointegration vector β_1 and by examining whether the resulting linear combination is stationary. These restrictions are defined by design matrices H , which are of the form $(1,0,0,0)$, $(0,1,0,0)$, etc. Each variable is tested separately by restricting the coefficients respectively in the H matrix. The null hypothesis for the Johansen test is stationarity.

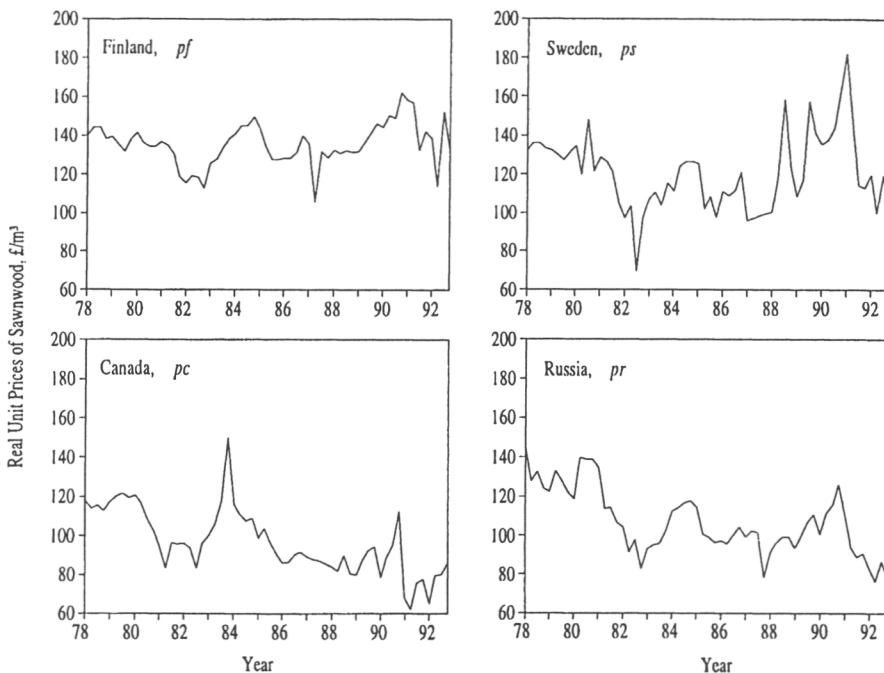
The hypothesis presented in Equation (4), i.e., the law of one price by pairs of prices, can be tested using the following restriction of Johansen:

$$\beta = H\varphi \text{ or } \Pi = \alpha\varphi' H \quad (A4)$$

where $H(p \times s)$, $\varphi(s \times r)$ and $r \leq s \leq p$

In (A4) the restrictions are defined by φ , which is an $s \times r$ matrix, where $r(p - s)$ is the number of restrictions and r is the number of cointegration vectors. The testable null hypothesis (4) between two price variables can be expressed as $\beta_{i1} = -\beta_{j1}$. With four variables, under $r = 1$, the design matrices, H , are of the form $(1, -1, 0, 0)$, $(1, 0, -1, 0)$, etc. If the restriction is rejected, the law of one price between the two prices is also rejected.

Appendix 2. Real Unit Prices of Soft Sawnwood Imports to the United Kingdom from Finland, Sweden, Canada, and Russia.



Testing arbitrage in newsprint imports to United Kingdom and Germany

Riitta Hänninen, Anne Toppinen, and Pertti Ruuska

Abstract: The arbitrage was tested for United Kingdom and German newsprint imports from Finland, Sweden, and Canada using the hypothesis of "the law of one price" (LOP). The LOP states that prices of homogenous commodities, defined in the common currency, are equal throughout the world. Concept of cointegration and Johansen's multivariate method were used in testing the LOP in the long run. The empirical data were quarterly covering the period 1980–1994. The hypothesis that the LOP holds for all three supplier countries simultaneously was not supported by the multivariate test results in either importing country. Cointegration tests rejected the simultaneous LOP between the three import prices in both markets. The LOP was also tested pairwise, and it was found to hold between Swedish and Canadian prices in the U.K. market.

Résumé : On a vérifié l'arbitrage en utilisant l'hypothèse de la « loi d'un seul prix » (LOP) pour les importations de papier journal du Royaume Uni et de l'Allemagne en provenance de Finlande, de Suède et du Canada. La LOP stipule que les prix des biens homogènes de consommation courante, définis par une devise commune, sont égaux à travers le monde. Le concept de la co-intégration et la méthode multivariée de Johansen ont été employés pour vérifier la LOP à long terme. Des données empiriques trimestrielles couvrant la période 1980–1994 ont été utilisées. L'hypothèse selon laquelle la LOP est valide pour les trois pays fournisseurs simultanément, n'a pas été confirmée par les résultats de la méthode multivariée pour aucun des pays importateurs. Les tests de co-intégration ont rejeté la simultanéité de la LOP entre les trois prix d'importations pour les deux marchés. La LOP a aussi été vérifiée par paire, et on a constaté qu'elle était validée entre les prix suédois et canadien sur le marché du Royaume Uni.

[Traduit par la Rédaction]

Introduction

The arbitrage implies that the prices of a particular commodity in various international markets are equalized when expressed in the same currency and consequently, the "law of one price" (LOP) holds. The LOP assumption implies perfect competition in the market, and it is a commonly used assumption in modelling commodity trade. The present study examines for arbitrage in the U.K. and German newsprint imports from Finland, Sweden, and Canada by using the multivariate cointegration approach. The focus of this study is on the long-run relationship between the prices of the three supplier countries.

A number of studies that model forest products markets and trade have assumed perfect competition, but there have also been several studies that have assumed imperfect competition, allowing for price differences across markets. The trade models assuming imperfect competition and allowing for price differences between products of different countries are commonly based on the Armington (1969) model. The basic assumption of the Armington (1969) model is that a product originating from a supplier country is not a perfect substitute for the same product originating from another supplier country. For example, price differences are assumed in studies that model trade in plywood (e.g., Chou and Buongiorno 1983), pulp and paper (e.g., Blatner 1989 and Laaksonen et al. 1997),

and sawnwood (e.g., Hseu and Buongiorno 1993; Hänninen 1994). The other approach, which assumes perfect competition and the LOP, is used, for example, in the Timber Assessment Market Model (Adams and Haynes 1980) and in the Global Trade Model (Dykstra and Kallio 1987). Brännlund et al. (1982) applied the assumption in modelling Sweden's trade in newsprint and the other forest products.

For the newsprint market, both perfect and imperfect competition have been assumed, and it is not clear which one is preferable in modeling this market. Previous models of the newsprint industry in North America (e.g., Dagenais 1976; Booth et al. 1991) were based on oligopolistic market structure with dominant firm price leadership, whereas Buongiorno and Gillies (1984) modelled newsprint trade, and Andersen (1993) studied the structure and development of the European newsprint market assuming competitive markets.

If empirical study assumes LOP when it actually does not hold, inferences are invalid and the forecasts are unreliable. Under perfect competition, relative costs determine export quantities, whereas under imperfect competition relative prices are the key determinants of exports. Policy recommendations based on competitive market model will not have the expected effects in the presence of imperfect competition, and the costs of incorrect policies may be large. Thus, in modelling and forecasting the demand for and trade of forest products, it would be useful to have direct tests for the existence of the LOP.

The present study tests the LOP in the U.K. and German newsprint markets between the main supplier countries, Sweden, Finland, and Canada by using Johansen's multivariate cointegration analysis (Johansen 1995). Earlier, Johansen's method in testing the LOP in forest products' markets has been used by, for example, Hänninen (1998), who tested the LOP for the U.K. sawnwood imports from the four main supplier countries.

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The LOP between the suppliers was rejected in the simultaneous and pairwise tests. On the other hand, Jung and Doroodian (1994), who studied the LOP in connection with four regional softwood lumber markets in the United States accepted the LOP. They found one cointegration vector, but did not determine for which one of the four prices does the LOP hold. One cointegration vector in the case of four prices means that the law cannot hold for all four prices simultaneously (e.g., Goodwin and Grennes 1994). The main finding of the present study is that simultaneous LOP was rejected in the U.K. and German markets. In pairwise analysis, the only case where the LOP was accepted was between the prices of Sweden and Canada in the U.K. market.

The paper is organized as follows. The outline for the testing procedure for the LOP is described in the next section, following by the data section and empirical results. Summary and conclusions are presented in the final section.

The LOP testing procedure

At the most aggregate level, long-run relations of price indices of total trade for different countries are examined using the theory of purchasing power parity (PPP). The LOP, being generally concerned with the commodity level, can be seen as a disaggregated counterpart of the PPP. Because of the importance of the LOP assumption it has been widely tested at the aggregate level of economy (e.g., Isard 1977; Ardeni 1989; Knetter 1993) using the concept of cointegration and the Engle and Granger (1987) method with two variables. The results from the previous studies generally do not support the LOP (see, for example, the survey of Officer 1986). For forest products there are only a few studies that examine the validity of the LOP.

Buongiorno and Uusivuori (1992) tested price differences in U.S. pulp and paper exports to different countries using bivariate cointegration analysis and found evidence that the LOP holds. However, when pairs of prices are considered separately, any linkages that may operate through a third market price are ignored. Also, the approaches using the Dickey–Fuller type methods actually test the null hypothesis that the LOP does not hold and thus perform the test indirectly. Our testing procedure differs from most of the earlier studies with respect to estimation method. We employ the multivariate cointegration method as developed by Johansen (1988) and Johansen and Juselius (1990). In using the multivariate method, the price equations can be estimated simultaneously. In Johansen’s procedure, it is possible to test directly the null hypothesis, that the LOP holds simultaneously in a multivariate context.

Traditionally, tests of the LOP have used the following bivariate procedure that we consider here as an example. Let p_{it} and p_{jt} be newsprint prices of two supplier countries, expressed in the currency of the destination country. Assuming no transportation costs, the LOP means that in equilibrium $p_{it} = p_{jt}$. We write the following regression equation, with prices in logarithmic form:

$$[1] \quad p_{it} = \alpha + \beta p_{jt} + u_t$$

If the LOP holds in [1], the restrictions $\alpha = 0$ and $\beta = 1$ should be accepted. The test is, however, conditioned on the assumption that prices are not simultaneously determined and that

both price series are stationary, i.e., $I(0)$ processes. If one of the two prices is nonstationary (integrated) and the other is not, the LOP between them is automatically rejected. If both prices have the same order of integration and if they form a cointegrated system, the theory of cointegration (Engle and Granger 1987) can be applied to examine the law. Specifically in [1], if prices are nonstationary (i.e., $I(1)$), and if u_t is stationary (i.e., $I(0)$), then p_{it} and p_{jt} are cointegrated. In this case, the LOP is said to hold in the long run and the cointegration test can be transformed into a stationarity test of the difference between the two prices (Baffes 1991). Consequently, the law can be tested by taking the difference:

$$[2] \quad z_t = p_{it} - p_{jt}$$

and testing whether $z_t = I(0)$. Because the ordinary least squares estimate of β from [1] is biased in small samples (Stock 1987), [2] was preferred over [1] in many of the previous studies that test for cointegration. For example, Buongiorno and Uusivuori (1992) used the difference of two price variables to test cointegration and examine the strong LOP.²

In the case of three price variables, which in this study were the unit prices of newsprint from Finland, Sweden, and Canada, the LOP holds if the suppliers’ prices in the long-run equilibrium are equal in a destination country. This situation can be described by $p_{1t} = p_{2t} = p_{3t}$ in equilibrium. It can also be presented in the form of a null hypothesis such as

$$[3] \quad H_0: p_{1t} = p_{2t}, \text{ and } p_{2t} = p_{3t}$$

If the prices are nonstationary, $I(1)$ processes, the hypothesis [3] can be tested by taking the differences $p_{1t} - p_{2t} = z_{1t}$ and $p_{2t} - p_{3t} = z_{2t}$ and testing whether z_{1t} and z_{2t} are $I(0)$.

Under the assumption that the LOP holds in the market, the prices are jointly determined. This means that simultaneity should be taken into account in testing. The cointegration method of Engle and Granger (1987) may suffer from the simultaneous equation problem. Moreover, in the case of three variables, the number of possible cointegration vectors may be one or two. By using the method of Engle and Granger it is possible to obtain only one vector. In addition, testing for parameter restrictions is not possible. These problems are resolved in the present study by using the multivariate maximum likelihood procedure of Johansen (1988) and Johansen and Juselius (1990, 1992). The method is briefly summarized in the following paragraphs. A comprehensive description of the method can be found in Johansen (1995) and Banerjee et al. (1993).

The rank test in the Johansen procedure defines how many cointegration vectors, r , can be found in the data. The cointegration rank is determined by estimating a p -dimensional VAR(k) model system (e.g., Johansen 1995, p. 11):

$$[4] \quad x_t = A_1 x_{t-1} + \dots + A_k x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t = 1, \dots, T$$

where x_t is a vector of empirical variables, which are the newsprint import prices from Finland, Sweden, and Canada into the United Kingdom and Germany. Thus, the dimension (p) of the

² Buongiorno and Uusivuori (1992) also tested a weaker form of the LOP. This was done by testing the integration of u_t in eq. 1 without restricting α and β , respectively, to zero and one in ordinary least squares regression.

Table 1. Percentage shares of Finland, Sweden, and Canada from imports of newsprint to the United Kingdom and Germany, 1980–1994.

Period	United Kingdom			Germany		
	Finland	Sweden	Canada	Finland	Sweden	Canada
1980–1987	35	20	33	24	36	11
1988–1994	20	32	29	19	36	10
1980–1994	28	26	31	22	36	10

VAR model is three. Symbol μ is a vector of constant terms, D_t is a vector of seasonal dummies, k is the lag length ($k = 1, \dots, N$) and ε_t is a vector of error terms assumed to be NID(0, Ω). Equation 4 is reparameterized in error-correction form (e.g., Johansen 1995, p. 89) such as

$$[5] \quad \Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t = 1, \dots, T$$

where Δx_t is an $I(0)$ vector. In [5] the constant term, μ , can be restricted to cointegration space in the estimation if it is assumed that there is no linear trend in data. $\Gamma_1, \dots, \Gamma_{k-1}$ and $\Pi = -I + \Pi_1 + \dots + \Pi_k$ are coefficient matrices. Γ describes short-term dynamics of the process and Π is the matrix of long-run coefficients. The coefficient matrix Π can be decomposed into a matrix of loading vectors, α , and a matrix of cointegrating vectors, β , such that $\Pi = \alpha\beta'$. The cointegration vectors define the stationary linear combinations (long-run relations) of the variables in x_t . Further, Johansen (1988) formulates a likelihood-ratio-based trace test for testing cointegration rank, r , where the null hypothesis is $H_0: r = r_0$ against $r_0 < r \leq p$.

When the number of cointegration vectors, r , has been determined, it is possible to test hypotheses under r by imposing linear restrictions on the matrix of cointegration vectors, β , and loadings, α (Johansen and Juselius 1990, 1992). The tests for these linear restrictions are asymptotically χ^2 distributed. According to Goodwin and Grennes (1994), simultaneous LOP between prices requires $n - 1$ cointegration vectors, where n is the number of price variables. In our case ($n = 3$ in each market) this would mean that $r = 2$. Under cointegration a one-to-one relationship between prices is required for the simultaneous LOP to hold (Baffes 1991). This restriction is justified since we use CIF (cost, insurance, and freight) prices in each destination country, and they are neutral to the effect of freight costs unlike free on board prices. Subsequently, hypothesis [3], i.e., the simultaneous LOP, can be tested with the following restriction on the β matrix assuming that two cointegration vectors exist in the data:

$$[6] \quad \beta = H\varphi = \begin{bmatrix} 1 & 1 \\ -1 & 0 \\ 0 & -1 \end{bmatrix} \varphi$$

where H is a design matrix giving the restriction on the β matrix and φ is a (2×2) matrix of the parameters reduced by means of H . If restriction [6] is rejected, the law is also rejected for all prices simultaneously, in which case the second testable hypothesis is that the LOP holds between any two prices.

In the market where the LOP holds, some prices may still be driving forces behind the other prices. Finally, we test

whether any of the three newsprint import prices in the United Kingdom or Germany has such a property. This can be tested by the weak exogeneity test (e.g., Silvapulle and Jayasuriya 1994). The weak exogeneity can be tested by imposing restrictions on the loading vectors of the variables to equal zero in the cointegration space. If this restriction is accepted for a variable, it is weakly exogenous and consequently does not adjust to deviations from any equilibrium relation, as defined by the cointegration relation (Johansen 1992).

Data

The data used in this paper consist of unit prices of newsprint imported from Finland, Sweden, and Canada to the United Kingdom and Germany. The data are quarterly, seasonally unadjusted, and cover the period from 1980 to 1994.³ The import prices of newsprint from Finland, Sweden, and Canada to the importing countries were described by the average import unit values (£/ton for United Kingdom and DEM/ton for Germany) based on CIF figures. Data on quantities and values of newsprint (SITC 641.1) imports to the United Kingdom were taken from the Overseas Trade Statistics of the United Kingdom (CSO). The newsprint import quantities and values to Germany (SITC 4801 0010, including item 4801 650 for year 1980, item 4801 790 for years 1981–1987, and item 4801 0090 for 1988 onwards) were obtained from Aussenhandel nach Waren und Ländern (Spezialhandel). The unit price series were deflated by producer price indices (1985 = 100) for the United Kingdom and Germany, obtained from the Main Economic Indicators (OECD).

Average import unit values were used to describe prices because more disaggregated data on the quality composition of each supplier's newsprint were not available. This means that newsprint imported from the three suppliers was assumed to be homogenous enough to allow reliable testing of the LOP using unit values. Relatively speaking, newsprint is the most homogenous grade among the different paper products and is used for similar end uses, mainly for printing newspapers and less demanding advertising publications. For the three supplier countries at hand, the homogeneity assumption was also justified, because these countries have similar fibre raw material basis, i.e., coniferous wood.⁴

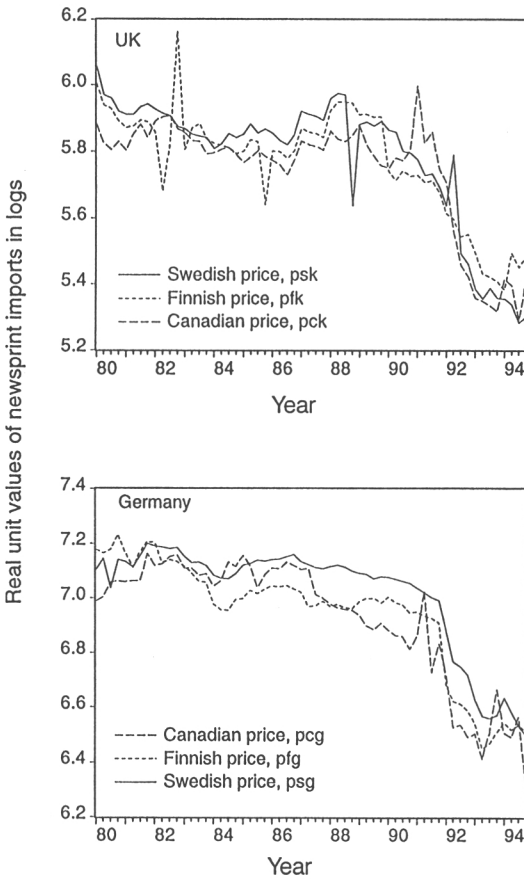
Empirical results

The three supplier countries have accounted for about 85% of newsprint imports to the United Kingdom and 68% to Germany during the period studied. Their average import shares are shown in Table 1. It can be seen that the combined import shares of these three countries have slightly fallen during the period, both in the United Kingdom and Germany. However, loss of import share has concerned only Finland and Canada. Sweden has managed to maintain its share in Germany and has

³ In principle it would be possible to obtain monthly price data. However, considering newsprint markets, quarterly frequency seems more suitable, since newsprint trade in Europe is often based on long-term contracts (see Andersen 1993).

⁴ Possible consequences due to differences in quality composition between different exporters or the effect of different ports of entry and exit, data errors, etc. cannot be taken into account, but they need to be borne in mind when interpreting empirical results.

Fig. 1. Real unit values of newsprint imports (in logs) from Finland, Sweden, and Canada to the United Kingdom and Germany, 1980 (first quarter) – 1994 (fourth quarter).



strongly increased it in the United Kingdom. In Germany, Sweden has been clearly the most significant exporter during the whole period. Canada’s average share has been only about 10% of total German imports. Germany is far less dependent on imported newsprint than the United Kingdom. The ratio of domestic newsprint production to imported newsprint has been around 90% in Germany and about 40% in the United Kingdom.

Graphs of newsprint prices are presented (real price series in logarithms) in Fig. 1. After stable price development during the 1980s, there has been a sharp fall in all newsprint prices in both importing countries because of economic recession in the early 1990s. In relative terms, newsprint imported from Sweden has had the highest, and the newsprint imported from Canada, the lowest average unit price in both the United Kingdom and Germany during the period studied.

Visual inspection of the graphs of the newsprint series (Fig. 1) suggests nonstationary prices. The test results from Augmented Dickey–Fuller (ADF) unit root test (Dickey and Fuller 1979) with one lag, presented in Table 2, indicate that

Table 2. Augmented Dickey–Fuller (ADF) unit root test results for prices of newsprint with one lag.

Price by supplier country	ADF test values by destination country	
	U.K.	Germany
Levels		
Sweden, ps	-0.46	-0.48
Finland, pf	-1.32	-1.36
Canada, pc	-2.22	-1.77
First differences		
Sweden, Δps	-7.21*	-5.73*
Finland, Δpf	-9.05*	-5.50*
Canada, Δpc	-4.83*	-7.94*

Note: *denotes the rejection of the null hypothesis of nonstationarity. The critical *t*-value is -3.49 at the 5% level.

all the levels of prices are nonstationary *I*(1) processes, while the first differences of prices are clearly stationary. ADF-test statistics were estimated from a model that includes a constant and a trend variable. We performed the test with 1 to 5 lags, and the results were found to be invariant to the use of different number of lags.

Hypothesis [3] was tested by determining the cointegration rank, i.e., the number of cointegration vectors, in both destination countries. To define the proper lag length for the cointegration estimation, model system [5] with alternative number of lags was fitted to the price data. Diagnostic tests for the residuals of each equation and the corresponding vector test statistics supported the model with four lags (*k* = 4) as a sufficient description of the data (Table 3). However, null hypothesis of normality was rejected for the residuals of all price equations for the United Kingdom and also for Finland’s price equation in Germany at the 1% level. According to Gonzalo (1994), non-normality of residuals does not bias the results for Johansen’s cointegration tests. Moreover, non-normality was mainly caused by outliers in the original data in U.K. customs statistics.

The cointegration test results from model [5] with four lags indicated one cointegration vector in the U.K. data, while no cointegration was found in the German data. The test results are presented with an unrestricted constant, allowing for linear trend in the data (see Fig. 1). According to the Johansen rank test, the hypothesis of no cointegration (*r* = 0) could be rejected and one cointegration vector was accepted at the 5% level in the United Kingdom (Table 4). In the case of the German price data, all the test statistics were well below the critical values and *r* = 0 could be rejected only at 50% level (cf. Johansen and Juselius 1990; critical value 20.19).

The three normalized eigenvectors (β) with weights (α) obtained from the cointegration estimation of model [5] for both destination countries are shown in Table 5. From the three vectors of the United Kingdom the first relation, β_1 , is interpreted as the cointegration vector, because the first relation is always most highly correlated with the stationary part of the process Δx_t , when corrected for the lagged values of the differences. The normalized α -coefficients are presented on the right-hand side of the long-term β -coefficients in Table 5. The α ’s can be interpreted as adjustment coefficients that describe

Table 3. Misspecification tests for residuals from Johansen's cointegration estimation of model [5] with four lags.

Equation	Tests for residuals and standard errors			Standard error (σ_e)
	Autocorrelation ^a ($F_{AR}(4,36)$)	Heteroskedasticity ^b ($F_{ARCH}(4,32)$)	Normality ^c ($\chi^2(2)$)	
United Kingdom				
Δpsk	0.80[0.53]	0.03[0.99]	27.86[0.00]*	0.07
Δpfk	0.54[0.71]	1.30[0.29]	36.91[0.00]*	0.08
Δpck	1.48[0.23]	0.10[0.98]	16.67[0.00]*	0.05
System	$VF_{AR}(36,77)=0.76[0.82]$		$V\chi^2(6)=89.49[0.00]*$	
Germany				
Δpsd	0.91[0.47]	0.42[0.79]	0.95[0.62]	0.02
Δpfd	0.33[0.86]	0.12[0.98]	9.66[0.01]*	0.04
Δpcd	2.05[0.11]	1.27[0.30]	5.02[0.82]	0.07
System	$VF_{AR}(36,77) = 1.80[0.02]$		$V\chi^2(6) = 16.98[0.01]*$	

Note: Δ denotes the first differences of the variables, psk, pfk, and pck denote the newsprint prices of Sweden, Finland, and Canada in the U.K. market, and psd, pfd, and pcd are the corresponding prices in the German market. Values in square brackets are marginal significance levels and * indicates that the null hypothesis is rejected at the 1% level.

^aAutocorrelation of the residuals was examined using the *F*-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged dependent variables.

^bHeteroskedasticity was tested using the *F*-form of the LM test against fourth-order autoregressive conditional heteroskedasticity.

^cNormality of the residuals was tested with the Doornik-Hansen test (Doornik and Hendry 1994). For further detail and test references, see Doornik and Hendry (1994).

Table 4. The result of the cointegration rank test.

Null hypothesis $H_0: r \leq i$	Eigenvalues (λ_i)	λ trace statistics ($T(i)$)	95% confidence values ($C(i)$)
United Kingdom			
$r=0$	0.35	35.33*	29.70
$r \leq 1$	0.18	11.40	15.40
$r \leq 2$	0.00	0.02	3.80
Germany			
$r=0$	0.28	23.09	29.70
$r \leq 1$	0.07	4.40	15.40
$r \leq 2$	0.00	0.16	3.80

Note: * indicates rejection of the hypothesis $r = 0$.

the average speed of adjustment of individual equations toward the estimated equilibrium state (see Johansen and Juselius 1990). A low coefficient indicates slow adjustment, while a high coefficient indicates rapid adjustment. For example, the coefficients of the α_1 -vector represent the weights with which Sweden's newsprint price enters the price equations of Finland and Canada.

The results of the rank test in Table 4 imply that hypothesis [3] of simultaneous holding of the LOP can be rejected for both destination countries, because $r < 2$. As there was no cointegration in the German market, the simultaneous long-run LOP cannot hold. In the U.K. market, the simultaneous LOP between three prices, i.e., hypothesis [3] from eq. 6, was rejected. The value of the $\chi^2(2)$ test statistic was 8.68, while the 5% critical level is 5.99. The rejection of hypothesis [3] indicates that the LOP cannot be valid for all three prices simultaneously. However, a stationary relationship between any two price variables is still possible in the United Kingdom. This is

Table 5. Normalized eigenvectors, β_1 , and their weights, α_1 , obtained from cointegration estimation of model [5].

Prices by countries	Eigenvectors			Weights		
	β_1	β_2	β_3	α_1	α_2	α_3
United Kingdom						
Sweden, psk	1.00	-0.51	-9.94	-0.23	0.39	0.00
Finland, pfk	-0.22	1.00	2.92	0.24	-0.35	0.00
Canada, pck	-0.96	-0.32	1.00	0.56	0.21	0.00
Germany						
Sweden, psd	1.00	-0.21	-2.29	-0.03	-0.05	-0.01
Finland, pfd	-1.40	1.00	1.02	0.10	-0.05	-0.01
Canada, pcd	-0.20	-0.99	1.00	-0.03	0.07	-0.04

why the law was tested in the next phase also for the pairs of prices under one cointegrating vector indicated by the trace test.

For a pairwise testing of the LOP, the coefficients of the cointegration vector, β_1 , were restricted. The long-term equilibrium relation, β_1 , normalized to Swedish price can be presented in vector form as

$$[7] \quad \beta_1 = [1.00 \quad -0.22 \quad -0.96]$$

where 1.00 is the coefficient of Swedish price, -0.22 is the coefficient for the Finnish price, and -0.96 is the coefficient for the Canadian price. The coefficients indicate that the LOP may hold between Sweden and Canada, but not between Finland and the others. The formal test was performed by restricting two price coefficients at a time to equal 1 and -1 and the third to equal zero in β_1 vector in [7]. Test results for the United Kingdom under one cointegrating vector, presented in Table 6,

Table 6. Test results for the LOP by pairs of prices obtained from the U.K. model by restricting β_1 under $r = 1$.

Testable prices	Hypothesis ($H_0: P_{it}=P_{it}$)	Test (LR test statistic ^a)
psk, pfk	$\beta_{11=1}, \beta_{21=-1}, \beta_{31=0}$	16.68*
psk, pck	$\beta_{11=1}, \beta_{21=0}, \beta_{31=-1}$	5.36
pfk, pck	$\beta_{11=0}, \beta_{21=1}, \beta_{31=-1}$	9.88*

Note: psk, pfk and pck are the newsprint prices of Sweden, Finland and Canada in the UK.
^aCritical χ^2 value is 5.99 at the 5% level. *denotes the rejection of the law of one price.

Table 7. Test results for weak exogeneity of supplier countries' newsprint prices in the United Kingdom.

Testable price	Test (under $r=1$) $\alpha_r=0$ (LR test statistic ^a)
Sweden, psk	1.32
Finland, pfk	1.30
Canada, pck	9.71*

^aCritical χ^2 value 3.84 at the 5% level.
 *denotes the rejection of weak exogeneity.

rejected the law for the price pairs for Sweden and Finland (psk, pfk) and Finland and Canada (pfk, pck). Only the restriction [1, 0, -1] was not rejected, which means that for the prices of Sweden and Canada (psk, pck), the test supported the LOP.

The results for the weak exogeneity test in Table 7 reveal that the Swedish and Finnish newsprint prices are weakly exogenous to the long-run parameters in the U.K. market. Thus, Swedish and Finnish prices do not adjust to deviations from any equilibrium state defined by the cointegration relation, but are rather determined outside the model system. For the Canadian price, weak exogeneity was rejected, suggesting that Canadian price is affected by the Swedish and Finnish prices in the U.K. market. Furthermore, the existence of the LOP between Swedish and Canadian prices in United Kingdom can be interpreted so that Canadian price is the adjusting part and the Swedish price is the determining part of the relationship. The result of the weak exogeneity tests could also reflect the fact that Scandinavian producers are dominating the European newsprint market, while Canadian imports to the area are the residual of the market demand (see also Brännlund and Löfgren (1995) on the results of testing for cyclical dumping in the newsprint markets).

Summary and conclusions

In this paper the long-run LOP was tested for U.K. and German newsprint imports from Finland, Sweden, and Canada by using the theory of cointegration (Engle and Granger 1987). Johansen's (1995) multivariate cointegration method was applied in the empirical analysis. The test results were calculated using quarterly import unit values deflated by producer prices for the period 1980-1994 ($T = 60$).

The cointegration tests indicate that the LOP is not valid in either market. The results supported one cointegration vector in the U.K. data, while in the German data no cointegration was found. Also the results of the pairwise tests in the U.K. market indicated rejection of the LOP except for a pair of Swedish and Canadian prices. Differences in cointegration rank test results between the two destination countries may be due to fact that Germany is less dependent on the imported newsprint than the United Kingdom. Therefore the price of their own newsprint production in Germany, which was not available, may affect the price structure of the German market. However, the exclusion of German price can be justified by the fact that newsprint production in Germany is mainly based on the use of recycled fibre (Andersen 1993). Therefore its price is not perfectly com-

parable with the prices of the mainly fibre-based newsprint imported from Sweden, Finland, and Canada.

The results of this study differ from the traditional assumption of trade theory, i.e., that commodity prices are perfectly arbitrated (e.g., Buongiorno and Gilless 1984). They also differ from the test results of Buongiorno and Uusivuori (1992), who accepted the LOP for U.S. paper exports across destinations. The results do agree, for example, with the results of Isard (1977), who compared export prices of different products between the United States and Germany; Knetter (1993), who found price differences between exporters for agricultural products; and Hänninen (1998), who tested the law for U.K. sawnwood imports across the main supplier countries.

Newsprint is a relatively homogenous product, so the arbitrage could be expected to hold. However, our results indicate long-run deviations of the LOP and support imperfect competition in the newsprint markets. Thus, the models allowing for price differences between the main supplier countries can be used in modeling the U.K. and German newsprint markets. The results of the study also stress the importance of formal testing for competition in forest products markets, which is often missing in the studies of forest products trade.

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IV

Exchange Rate Changes and the Finnish Sawnwood Demand and Price in the UK Market

Riitta H. Hänninen

Hänninen, R.H. 1998. Exchange rate changes and the Finnish sawnwood demand and price in the UK market. *Silva Fennica* 32(1): 61–73.

This paper examines the long-run influence of exchange rate changes on the Finnish sawnwood price in the United Kingdom (UK) using quarterly data for the period 1978–1994. The degree of influence was measured by a pass-through coefficient (PT) obtained from a markup pricing relation of a system model. The model, which included export demand and price equations, was estimated with the cointegration method of Johansen. The results indicated a large PT, which means that exchange rate changes are reflected almost proportionately in Finnish export price expressed in pounds sterling. Thus, the Finnish price of sawnwood in pounds has lowered as a result of depreciation of the Finnish markka (FIM). This has improved Finnish competitiveness and market share in the UK. Appreciation of the FIM has had the opposite effect. It seems that Finnish exporters have made use of depreciations and devaluations of the FIM to maintain and increase their market shares but not necessarily their markups. For Finland, which is in the process of joining the European economic and monetary union (EMU), knowing the size of the PT is also important in assessing the economic impact of membership.

Keywords exchange rate, export demand, Johansen's cointegration method, pass-through, sawnwood price, United Kingdom

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1 Introduction

This study examines the effects of exchange rate changes on Finland's sawnwood exports to the United Kingdom (UK), which is a major market for Finnish sawnwood. In 1996 the UK account-

ed for 18 percent of Finland's sawnwood exports (Facts and Figures 1997). Exchange rate effects are analysed by a demand and supply model for Finnish sawnwood exports. The focus of the study is on the exchange rate pass-through (PT), i.e. the percentage change in export prices

associated with a one percent change in the exchange rate.

If the exchange rate has a small effect on foreign currency prices, currency realignments should not have much effect on exports. Previous results are scanty for forest products, but they indicate that trade flows of forest products (measured in quantities) have not been very sensitive to exchange rate variations (Buongiorno et al. 1988). Also Uusivuori and Buongiorno (1990) found no long-run relationship between exchange rates and quantities of paper exports from Finland and Sweden to the USA. For Finland, which is in the process of joining the EMU, the magnitude of the exchange rate effect is important in assessing the economic impact of membership. In the EMU, it will no longer be possible to improve competitiveness by manipulating the exchange rate. Estimates of the degree of pass-through are also informative as to market competition, which is important in modeling trade flows.

The pass-through concept can be defined as the extent to which a change in a country's nominal exchange rate induces a price change in terms of the foreign currency. The magnitude of the pass-through also reflects the extent of market competition. For example, in perfect competition, export prices in foreign currency do not change as a result of a devaluation (revaluation) of the exporter's currency. In this case, Finnish exporters' prices in FIM would increase (decrease), their markup margins would increase (decrease) and the PT would be zero. On the other hand, in imperfect competition Finnish exporters could change their foreign currency prices when the exchange rate changes. If exchange rate fluctuations were partly reflected in foreign currency prices and partly in FIM prices, PT would be between zero and unity. If PT is unity, Finnish exporters lower (raise) their foreign currency export prices pro rata to a devaluation (revaluation) of the FIM

Pass-through has not been extensively studied for forest products trade, although it has been analyzed for trade in many other commodities and for aggregate imports and exports (see e.g. the survey Menon 1995). PT is found to vary widely by industry (e.g. Feenstra et al. 1996), which means that it is necessary to study pass-through at the commodity level. Most of the few

studies analyzing the PT for forest products deal with the USA. For example, there are very few results for Finnish exports. In the study of Vesala (1992) PT for export prices of Finnish paper manufactures was found to be between 0.66 and 0.69 for western Europe and between 0.16 and 0.30 for the USA. According to Vesala, the smaller PT for the USA is due to the large US domestic market, in which the Finnish share is much smaller than in Europe. Other studies have also found that prices of several commodities imported to the USA only slightly reflect changes in exchange rates (e.g. Yang 1997).

Uusivuori and Buongiorno (1991) estimated PT for US forest products exports to Europe and Japan. Pass-through was incomplete in most of the product categories. However, in lumber exports to Japan, PT was high: from 0.79 to 1.04, depending on the species. Also in the study of Menon (1993a), who estimated PTs for Australian imports, wood products had a relatively high PT (0.80), while for paper and board the PT was 0.45. The earlier results indicate that PT is higher for wood industry products than for paper products. However, exact conclusions about the size of the PT for forest products are difficult to draw, because studies are scarce and they concern only a few products and countries.

Based on the earlier models for Finnish sawnwood exports to the UK (e.g. Hänninen 1986, Tervo et al. 1988, Hänninen 1994) and studies that test arbitrage in UK sawnwood imports (Hänninen 1998), it can be assumed that the PT is relatively large. The assumption is also supported by the structure of UK sawnwood imports, which is dominated by four large supplier countries, i.e. Finland, Sweden, Canada and Russia.

The analysis of the present study is based on a structural multivariate model formulated for Finnish sawnwood exports to the UK. The model consists of an export demand equation and a markup price equation. In earlier studies simultaneous multivariate models have not been applied in analyzing pass-through in forest products markets and they have rarely been applied in other commodities markets (see however Rockerbie 1992, Menon 1993a, 1993b, Kongsted 1996). Previously used models have usually been bivariate, with a price variable being regressed on an exchange rate (e.g. Knetter 1989,

1993, Pick and Park 1991, Uusivuori and Buongiorno 1991).

Multivariate approaches based on the markup concept have also been applied, but these applications have been single equation models (e.g. Dornbusch 1987, Hooper and Mann 1989, Athukolara 1991, Athukolara and Menon 1994, 1995, Hung et al. 1993). The present study uses the Johansen (Johansen 1988, Johansen and Juselius 1990 and 1992) multivariate cointegration method for nonstationary data for estimation purposes. This allows for the estimation of a model system, unlike the traditional Engle and Granger (1987) procedure.

The findings indicate that PT is large, which means that in the long-run changes in the exchange rate (FIM/pound sterling) are almost completely reflected in Finnish prices as measured in pounds. Thus, for example, depreciations and devaluations of the FIM have improved Finnish competitiveness and market share by lowering the relative Finnish price in the UK market.

2 Model

In the present study, a theoretical model is formed to describe Finnish sawwood exports to the UK. The model includes a demand equation and a price equation. An estimate for the exchange rate pass-through (PT) is obtained from the price equation. The same type of model construction has earlier been applied by Kongsted (1996) in modeling Danish manufacturing exports.

The model is constructed on the assumption that competition between supplier countries is imperfect in UK sawwood imports. Thus, the elasticity of substitution model describing demand for goods from different origins is used to describe the export demand for Finnish sawwood. The derivation of the export demand assumes a two-stage optimization of a representative sawwood importer in the UK. First, costs are minimized subject to the expenditure on a good (e.g. sawwood) and, secondly, this expenditure is allocated optimally between the products from different countries of origin. Thus, the export demand for Finnish sawwood can be represented as

$$X_f = b_f \eta X_o (P_f/P_o ER)^\eta \quad (1)$$

where X_f and P_f are the Finnish quantity (1000 m³) and nominal unit price of sawwood exports (FIM/m³) to the UK market, X_o and P_o are the respective quantity and price (£/m³) of competitors' sawwood, ER is the nominal exchange rate (FIM/£), b_f is a constant and η is the elasticity of substitution (assumed to be constant). After logarithmic transformation of the variables in (1), the export demand relation becomes

$$x_f = -\eta (p_f - p_o - er) + x_o + c + \varepsilon \quad (2)$$

where lower-case letters denote logarithmic values of the corresponding upper-case letters in equation (1), c is a constant term and ε is a disturbance term. The symbol η is the price elasticity of demand and equation (2) is homogeneous of degree zero in the nominal variables. In this elasticity of substitution model (2), the change in the relative quantities demanded is assumed to be proportional to the change in the relative price of exports.

An equation for Finnish sawwood price is based on a markup model that has earlier been applied in the estimation of pass-through e.g. by Dornbusch (1987), Hooper and Mann (1989), Athukolara (1991), Athukolara and Menon (1995), Hung et al. (1993). In deriving the price equation, it is assumed that a representative Finnish exporter firm produces exclusively for an imperfectly competitive UK market, employs constant-returns-to-scale technology and unit production cost, C_f . The firm maximizes profit by taking the competitors' price and the supply of competitors' sawwood as given and by setting the price in FIM, P_f , as a constant markup over unit production costs, C_f . With X_f denoting export quantity, the exporter's profit, V_f , is defined as

$$V_f = (P_f - C_f) X_f \quad (3)$$

Profit maximization yields

$$P_f = C_f \eta / (\eta - 1) \quad (4)$$

where η is the price elasticity of demand. According to Hung et al. (1993), a more general

case, in which competitors' prices determine the exporter's price, can be presented by using the concept of a variable markup. Then, a variable markup can be defined by assuming that the coefficient η depends partly on price competitiveness in the export market. Competitiveness can be described as the relative price $(P_o ER)/P_f$, where $(P_o ER)$ is the competitors' price in the terms of exporter's currency ($ER = \text{FIM}/\text{£}$). Thus, the price elasticity of demand is

$$\eta = \eta((P_o ER)/P_f) \quad (5)$$

From (4) and (5), the pricing behavior of a profit maximizing exporter can be described as a variable markup over the unit cost:

$$P_f = \phi C_f \quad (6)$$

The variable markup, ϕ , depends on the relative price and may be approximated as

$$\phi = \phi((P_o ER)/P_f) = \phi'((P_o ER)/P_f)^\theta \quad (7)$$

where $\theta (\geq 0)$ is the relative price elasticity of the markup. The constant markup is obtained if $\theta = 0$ and $\phi' = \eta/(\eta - 1)$. The second equality in (7) comes from the log-linear approximation of the nonlinear function ϕ . Substituting (7) into (6) and taking a logarithmic transformation, a relation for the price of Finnish sawnwood in the UK market is obtained:

$$p_f = \delta + (1 - \gamma)(er + p_o) + \gamma c_f + u \quad (8)$$

where $\gamma = 1/(1 + \theta)$, $0 < \gamma \leq 1$, $\delta = \ln \phi'/(1 + \theta)$ is a constant, and u is a disturbance term that captures all other factors. The other symbols are the same as above. Lower-case variables denote logs of the corresponding upper-case variables. The export price, p_f , is homogenous of degree zero in the exchange rate and competitors' prices, and the equality restriction is imposed on the coefficients of p_o and er in the estimation.

The degree of exchange rate pass-through (PT) can be derived as the absolute value of the exchange rate elasticity of export price measured in foreign currency (e.g. Kongsted 1996). From equation (8) we obtain

$$PT = -(\partial(p_f - er)/\partial er) = \gamma, \quad 0 < PT \leq 1 \quad (9)$$

If $\gamma = 0$, then the exchange rate affects the Finnish price in FIM (equation 8) and thus the changes in er (FIM/£) are absorbed by the variable markup. This means that Finnish exporters do not pass through changes in the exchange rate to their export prices in pounds, i.e. $PT = 0$, which indicates perfect competition and the existence of the law of one price in the market.

If $\gamma = 1$, Finnish export prices in FIM, p_f , are proportional to production costs, c_f , and $PT = 1$. This indicates that Finnish exporters fully pass through er changes to their export prices in pounds and keep the markup constant. This indicates imperfect competition in the market. Between the above extremes ($0 < \gamma \leq 1$) we have incomplete pass-through. In this case, changes in er are reflected partly in FIM-denominated and partly in foreign currency-denominated export prices.

The model used in the study may naturally be affected also by factors other than those included in the model. If production costs include an important imported component, it is evident that the degree of pass-through will be less than one, even if $\gamma = 1$. In the Finnish sawnwood industry, inputs are for the most part of domestic origin, which suggests that this effect is not very important.

3 Estimation of the Model

3.1 The Data

During the period studied, the FIM fluctuated widely with respect to the pound sterling (graph 1C in Appendix 1). Moreover, the exchange rate regime was revised several times. The FIM was fixed with respect to other currencies in the years 1978–1992. However, the value of the FIM was changed by specific decisions of the Finnish government, being revalued in 1979, 1980 and 1989 and devalued in 1978, 1982, 1991 and 1992. Revaluations can be seen as declines and devaluations as rises in the graph (1C). During the fixed exchange rate period, the value of the FIM was measured by a currency index that included

the currencies of Finland's most important trading partners, until June 1991. After this, the FIM was unilaterally linked to the ECU basket until September 1992 and then allowed to float until October 1996.

The development of the exchange rate (graph 1C) can be divided into three subperiods. The period of depreciation, 1978–1980, was followed by a long period, 1981–1990, during which the FIM appreciated by about 22 per cent. In the years 1990–93 the FIM again depreciated by about 26 per cent with respect to the pound sterling. The same subperiods can also be distinguished in the development of the other variables. A comparison of the graphs (1B, 1C, 1D, 1F) indicates that FIM depreciations have lowered the relative Finnish price in pounds and increased export quantities and Finnish market share in the UK. FIM appreciations have had the opposite effects.

It seems that exchange rate changes have affected Finnish competitiveness and market share by changing the Finnish price in terms of pounds, while the markup has remained relatively constant. Comparison of the markup on production costs ($p_f - c_f$) and the exchange rate, er , indicates that exchange rate fluctuations have not had much of an impact on the markup. The markup series is relatively stationary up to 1990. After 1990, when FIM was devalued two times by large percentages, the markup jumped up to a higher level (graph 1E). It is possible that the markup absorbed a larger part of the exchange rate change in the 1990s than in the 1980s. If this is true, PT has decreased in the 1990s. It is also likely that the markup increased partly as a result of a decrease in production costs. Finnish prices of sawlogs decreased by about 30 percent due to the recession in the Finnish economy in the years 1990–93. Exchange rates and prices of products and inputs seem to have been much more turbulent in the 1990s than earlier, which makes it difficult to draw conclusions.

The data used in this study were seasonally unadjusted, quarterly, and covered the period from 1978 to 1994. Finnish and competitors' prices of sawnwood were described by import unit values of sawnwood to the UK (£/m³) based on CIF figures (CIF includes cost, insurance and freight). The Finnish price in pounds was con-

verted into FIM, p_f , at the nominal exchange rate. The competitors' price, p_o , is a (quantity-share) weighted unit price of imports (£/m³) from Sweden, Canada and Russia. Competitors' import quantity, x_o , is the sum of UK imports from these three countries. Data on quantities and values of sawnwood (SITC 248.2-3) imports were taken from the Overseas Trade Statistics of the United Kingdom (CSO) for the years 1978–1990 and from the intra- and extra-EU trade statistics (European Commission) for the rest of the observations.

The exchange rate variable, er (FIM/£), was the nominal quarterly average. Price observations for the period 1990:1–1994:4 were transformed using the £/ECU rate. Both exchange rates were obtained from International Monetary Fund (IMF) statistics. Production costs of Finnish sawnwood, c_f , were described by the production price index (1990 = 100, manufacture of wood and furniture), which was obtained from the Bulletin of Statistics (Statistics Finland). The use of the production price index as a proxy for the domestic cost of the sawnwood industry may cause uncertainty in the testing of PT, but it is used because proper cost variables are not available. Another possibility would have been to describe the production costs by the price of rawmaterial. Quarterly time series for prices of sawlogs are however available only from year 1985, which would make the estimation period too short.

3.2 Estimation Method

The cointegration estimation of the model comprising Equations (2) and (8) involves a vector of six time series. Provided that the data is cointegrated, it may be possible to identify two cointegration vectors the coefficients of which describe the long-run equilibrium relationships implied by equations (2) and (8).

The present study uses Johansen's cointegration method (Johansen 1988, 1995), which is suitable for the estimation of nonstationary data. The main advantage of Johansen's method compared to the earlier alternative multivariate methods (e.g. Engle and Granger 1987) is that it makes it possible to estimate and identify multiple

cointegration vectors. In more complex cases than the two-variable case, there may exist up to $p - 1$ (p denotes the number of variables) cointegration vectors. The earlier methods do not solve this problem. When a cointegration relationship was identified, it was assumed to be unique. This need not be true in a multivariate case. Johansen's method provides a log-likelihood ratio test statistic for determining the number of cointegration vectors in the data.

If there exist more than one cointegration vector linking the variables together, the relationship estimated by the Engle-Granger method will be invalid. The relationship may simply represent complex linear combinations of all the cointegration vectors. Johansen's method, unlike the Engle-Granger procedure, also accommodates short-run dynamics in the cointegration regression. This helps to reduce biases and improves efficiency in using the information content of the data in the estimation. Johansen's method also enables testing of several economic hypotheses by means of linear restrictions in the same cointegration framework.

The stationarity of the variables was not tested separately prior to the analysis, as would be the case e.g. if the Engle and Granger (1987) method had been used. The property that unit vectors can be cointegration vectors in Johansen's model means that one can include in the cointegration analysis $I(1)$ or $I(0)$ variables provided they are economically meaningful (Johansen 1995, p. 74). By including a stationary variable in the vector of variables, X , one adds an extra cointegration vector, i.e. an extra dimension in cointegration space.

The analysis of the present study starts by estimating an unrestricted statistical VAR model including all six variables. The method of Johansen (1988) and Johansen and Juselius (1990) uses a statistical model that is a p -dimensional VAR process of order k , where p is the number of variables. The VAR model can be formulated as

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t = 1, \dots, T \quad (10)$$

where Δx_t is an $I(0)$ vector of the six first-differenced variables. In the VAR model the constant

term, μ , can be restricted to cointegration space so as to represent the absence of a linear trend in the data. However, in the present study a linear time trend was assumed to exist in the data and thus the constant term was not restricted. D_t is a seasonal dummy and k is the lag length. Introducing a sufficient number of lags usually produces a well-behaved error term $NID(0, \Omega)$. $\Gamma_1, \dots, \Gamma_{k-1}$ and $\Pi = -I + \Pi_1 + \Pi_2 + \dots + \Pi_k$ are coefficient matrices and ε_t is a vector of error terms assumed to be normally and independently distributed with expectation zero and variance matrix Ω .

In (10) the level terms capture the long-run steady state relationships and the first-difference terms show the short-run dynamics. Π is the matrix of long-run coefficients of the lagged levels. It can be decomposed into a matrix of loadings, α , and a matrix of cointegrating vectors, β , i.e. $\Pi = \alpha\beta'$. If x_t is $I(1)$, i.e. integrated of order one, the components of x_t are cointegrated. The number of cointegration vectors that exist among the variables is determined by estimating the rank of the matrix Π . Johansen's method (Johansen 1988) formulates two likelihood ratio tests for the cointegration rank, r , from which the trace test is applied here. The trace test for testing that there are at most r cointegration vectors in the set of p variables is defined as

$$\text{Trace}(r) = -T \sum_{i=r+1}^p \ln(1 - \lambda_i) \quad (11)$$

where T is the number of observations and the λ_i 's are the smallest squared canonical correlations (eigenvalues). When the number of cointegration relations, r , is determined it is possible to test hypotheses on the long-run matrix, $\Pi = \alpha\beta'$. The present study tests restrictions on the β_{ij} coefficients. In the Johansen framework these tests are conducted under the hypothesis

$$\beta = H\varphi \text{ or } \Pi = \alpha\varphi'H' \quad (12)$$

where H is a $(p \times s)$ matrix, φ is a $(s \times r)$ matrix and $r \leq s \leq p$. Restrictions, s , are defined by φ , and r is the number of cointegration vectors. The exclusion test (excluding a variable from a cointegration relation) was performed in order to iden-

tify the economic long-run relation as being that represented in equations (2) and (8). The testable null hypothesis for exclusion of a variable was $\beta_{ij} = 0$. Homogeneity between Finnish export price, p_f , and the other nominal variables was tested by restricting the respective coefficients accordingly. Finally, the markup relationship was tested by examining if the coefficient of c_f could be restricted to unity.

4 Results

4.1 Cointegration of the Empirical Variables

The cointegration estimation was based on the VAR(4) model given by equation (10), with six equations ($p = 6$) for the period 1978–1994. The lag length, k , of the VAR model was determined by the Schwarz (SC) and Hannan-Quinn (HQ) information criteria, using likelihood ratio tests. Starting from $k = 5$ (see Doornik and Hendry 1994, p. 287), a reduction of the VAR from $k = 5$ to $k = 4$ was accepted. Because the reduction from $k = 4$ to $k = 3$ was rejected, $k = 4$ was used for further modeling.

The diagnostic tests on the residuals of the VAR(4) model are presented in Table 1, and they support the model with $k = 4$. Autocorrelation of the residuals was examined using the F-form of the Lagrange Multiplier (LM) test, which

is valid for systems with lagged dependent variables. The null hypothesis of no serial autocorrelation was accepted at the 5 percent level. Heteroskedasticity was tested using the F-form of the LM test against 4th order autoregressive conditional heteroskedasticity. The null hypothesis of no heteroskedasticity was accepted at the 5 percent level. Normality of the residuals was tested by means of the Doornik-Hansen test (Doornik and Hendry 1994) and the null hypothesis of normality was accepted for all the equations. Also the corresponding vector tests for the equation system accepted normality and indicated no autocorrelation. For further details and references concerning these tests, see Doornik and Hendry (1994).

The results of the cointegration estimation of the VAR(4) model indicate that $r = 2$ (Table 2). According to Johansen's trace test, the hypotheses of $r = 0$ and $r \leq 1$ can be rejected. Thus two cointegration vectors are accepted at the 5 percent level. The eigenvectors (β_j) and their weights (α_j) obtained from the cointegration estimation of model (10) are shown in Table 3. Of the six eigenvectors, the first two relations (β_1 and β_2) are most highly correlated with the stationary part of the process Δx_t corrected for the lagged values of the differences. Thus, β_1 and β_2 are the two cointegration vectors determined by the model (Johansen 1995). They are normalized by the coefficients of Finnish export quantity, x_f , and Finnish export price, p_f .

The normalized vectors of loadings, α_j , are

Table 1. Misspecification tests for the residuals of the VAR(4) model.

Equation	Tests for the residuals and the standard errors			
	Autocorrelation $F_{AR}(4,32)$	Heteroskedasticity $F_{ARCH}(4,28)$	Normality $\chi^2_{N(2)}$	Standard errors σ_e
$\Delta(x_f)$	0.97[0.44]	0.02[1.00]	0.05[0.98]	0.16
$\Delta(p_f)$	1.96[0.12]	0.33[0.86]	2.00[0.37]	0.05
$\Delta(er)$	0.44[0.78]	0.22[0.93]	0.99[0.61]	0.04
$\Delta(c_f)$	1.14[0.36]	0.48[0.75]	4.01[0.13]	0.01
$\Delta(p_o)$	2.14[0.10]	0.14[0.97]	0.77[0.68]	0.07
$\Delta(x_o)$	0.69[0.60]	0.34[0.85]	0.32[0.85]	0.12
System:	$VF_{AR}(144,48) = 1.10[0.36]$		$V\chi^2_N(12) = 9.98[0.61]$	

Note: Values in square brackets are marginal significance levels.

Table 2. Results for the cointegration rank test.

Null hypothesis	Eigenvalues	λ trace statistics	95% critical values
$H_0: r \leq i$	λ_i	T(i)	C(i)
$r = 0$.56	122.9 *	94.20
$r \leq 1$.36	70.83 *	68.50
$r \leq 2$.29	42.23	47.20
$r \leq 3$.22	20.69	29.70
$r \leq 4$.06	4.88	15.40
$r \leq 5$.02	1.18	3.80

Note: * indicate rejection of the null hypotheses implying that the rank is 2.

presented under the β_j vectors in Table 3. The α_{ij} 's represent the weights with which the error-correction terms enter each equation and they indicate the average speed of adjustment toward the estimated equilibrium state. A low coefficient indicates slow adjustment, while a high coefficient indicates rapid adjustment. Table 3 shows that the loadings of the two cointegration vectors (β_1 and β_2) are rather low.

4.2 Long-run Exchange Rate Pass-through

The two estimated cointegration vectors are tested with restrictions in order to identify the model system comprising export demand and price relations for Finnish sawnwood. The unrestricted cointegration vectors from Table 3 are

$$\beta_1: 1.00x_f - 6.75p_f - 4.49er + 8.90c_f - 0.89p_o - 2.30x_o \text{ and} \tag{13}$$

$$\beta_2: 0.44x_f + 1.00p_f + 0.57er - 0.72c_f - 1.26p_o + 0.50x_o \tag{14}$$

where, x_f and p_f are Finnish quantity and unit price (FIM/m³) of sawnwood exports to the UK and x_o and p_o are the respective quantity and price (£/m³) of competitors' sawnwood. er is the exchange rate (FIM/£) and c_f is the unit cost of Finnish sawnwood output.

The first cointegration vector, β_1 (equation 13), was identified, as was equation (2), by excluding Finnish production cost (c_f) from the relation and

Table 3. Normalized eigenvectors, β_j , with corresponding weights, α_j , obtained from the unrestricted cointegration estimation.

Variables	Eigenvectors					
	β_1	β_2	β_3	β_4	β_5	β_6
x_f	1.00	0.44	-0.03	0.83	-0.41	-0.81
p_f	-6.75	1.00	-0.48	1.50	-0.82	-0.25
er	-4.49	0.57	1.00	-2.35	0.97	2.29
c_f	8.90	-0.72	0.26	1.00	-0.22	0.21
p_o	-0.89	-1.26	0.38	-2.16	1.00	-0.32
x_o	-2.30	0.50	0.27	-1.15	0.06	1.00

Variables	Weights					
	α_1	α_2	α_3	α_4	α_5	α_6
x_f	-0.05	-0.09	-0.29	0.26	0.24	0.11
p_f	0.05	-0.03	0.16	-0.17	-0.10	0.01
er	-0.01	-0.02	-0.52	-0.06	-0.03	0.01
c_f	-0.01	0.04	0.05	-0.02	-0.01	0.00
p_o	-0.04	-0.04	0.57	0.14	-0.11	0.03
x_o	0.16	0.15	0.10	0.32	0.06	0.05

Symbols: x_f = Finnish quantity, p_f = Finnish price (FIM), er = exchange rate (FIM/£), c_f = Finnish unit cost, p_o = competitors' price in pounds sterling and x_o = competitors' quantity.

assuming that the coefficient of x_o is $\beta_{61} = -1$. The exclusion of c_f was accomplished by restricting its long-run coefficient, β_{41} , to zero. The second cointegration relation, β_2 (equation 14), was identified, as was equation (8), by excluding Finland's and the competitors' quantity.

When the above restrictions and the homogeneity assumption required by economic theory were tested for the demand and price relations, the test rejected this structure (structure I in Table 4). The homogeneity of the equations in the nominal variables p_f , c_f and ($p_o + er$) implies that the coefficients of p_o and er should be equal and that the coefficient of c_f should equal the difference between the coefficients of p_f and er in the cointegration vectors. However, the coefficients of the unrestricted equation (13) in particular indicate that homogeneity does not necessarily hold. Thus testing was continued by applying structure II (Table 4), where the homogeneity restriction is applied only in the price equation.

The structural form (II) is accepted. The type of demand equation (1) applied in the present

Table 4. Tests for the restrictions on the unrestricted cointegration vectors β_1 and β_2 under $r = 2$.

Variables/ LR-tests	Restricted and normalized demand (β_{11}) and price (β_{12}) relations ($i = 1, \dots, 6; j = 1, 2$)					
	(I)		(II)		(III)	
	$\beta_{3j} = \beta_{3j}$ and $\beta_{31} = -\beta_{21}$, $\beta_{41} = 0, \beta_{61} = -1, \beta_{12} = 0, \beta_{62} = 0, \beta_{42} = -(\beta_{22} + \beta_{32})$		$\beta_{31} \neq -\beta_{21}$, others unchanged		Mark-up: $\beta_{42} = 1$, others unchanged	
	β_{11}	β_{12}	β_{11}	β_{12}	β_{11}	β_{12}
x_f	1.00	0.00	1.00	0.00	1.00	0.00
p_f	2.13	1.00	1.57	1.00	2.44	1.00
er	-2.13	-0.04	-6.16	-0.08	-6.56	0.00
c_f	0.00	-0.96	0.00	-0.92	0.00	-1.00
p_o	-2.13	-0.04	-6.16	-0.08	-6.56	0.00
x_o	-1.00	0.00	-1.00	0.00	-1.00	0.00
LR	$\chi^2(1/4 4) = 14.62^*$		$\chi^2(1/4 3) = 6.84$		$\chi^2(1/4 5) = 7.63$	

Notes: 1) Symbols: x_f = Finnish quantity, p_f = Finnish price (FIM), er = exchange rate (FIM/£), c_f = Finnish unit cost, p_o = competitors' price in pounds sterling and x_o = competitors' quantity.
2) * indicates rejection of the restricted model structure.

study is not often tested for homogeneity in earlier applications. The homogeneity condition has usually been satisfied simply by expressing the exporter's and competitors' prices in a common currency in the form *PF/PO*. Because the price relation fulfills the restrictions and produces the PT estimate, the estimation is continued by restricting it further.

In the accepted structure (II), the price relation β_2 resembles the markup pricing relation with $\gamma = 1$ (Table 4). The coefficient of unit cost, c_f , is close to unity (-0.92), while the coefficients of the exchange rate (er) and competitors' price (p_o) are close to zero. Finally, the markup assumption was tested by restricting the price relation accordingly. The resulting structure (III) is also accepted by the test and the final long-run equilibrium relations can be presented as

$$\beta'_1: 1.00x_f = -2.44 p_f + 6.56(er + p_o) + 1.00x_o \text{ and} \tag{15}$$

$$\beta'_2: 1.00p_f = +0.00(er + p_o) + 1.00c_f \tag{16}$$

where the symbols are the same as above. Because the demand equation failed the homogeneity test, interpretation of its coefficients is problematic. However, the signs of the own-price and exchange rate elasticities of Finnish sawnwood export demand are consistent with the eco-

nomie theory. Moreover, the magnitude of the price elasticity (-2.44) is between the earlier estimated results of Hänninen 1994 (-1.71, estimated from annual data, 1976-90) and of Tervo et al. 1988 (-3.1, estimated with Almon polynomials, 1-12 lags from quarterly data, 1966-85). The relatively large elasticity of the exchange rate (6.56) implies a large effect of exchange rate on Finnish sawnwood exports to the UK.

The restricted price relation (16) representing markup pricing, with $\gamma = 1$, indicates that Finnish sawnwood export price in FIM is proportionate to production cost and the effect of the exchange rate on the FIM price is very small (zero). Thus, exchange rate pass-through is large. A large pass-through coefficient indicates that exchange rate changes are reflected almost pro rata in the Finnish export price in pounds sterling.

5 Conclusions

The study examined the long-run exchange rate pass-through (PT) for the Finnish price of sawnwood in the UK market by estimating a demand and price equation system. The data were quarterly and covered the years 1978-1994. Johansen's cointegration method, which is suitable for analyzing nonstationary data, was used in the

estimation. The main advantage of Johansen's method in this study is that it allows for the estimation and identification of a system model, unlike the traditional Engle and Granger (1987) procedure. It also accommodates short-run dynamics in the cointegration regression, which helps to reduce biases. Johansen's method also enables testing of several economic hypotheses by means of linear restrictions in the same cointegration framework.

According to Johansen's rank test, two cointegration vectors were determined among the variables. These vectors were restricted in order to identify them with the theoretical export demand and price relation. The restrictions were accepted in both cointegration vectors, but in the export demand relation (15) the assumption of price homogeneity did not hold. Thus, the coefficients of the demand equation are problematic to interpret. However, the price and exchange rate elasticities were of the expected signs and the magnitude of the price elasticity was consistent with the earlier estimated results (Hänninen 1994 and Tervo et al. 1988). The elasticity of the exchange rate estimated from the demand equation was relatively large, implying a large effect for the exchange rate on Finnish sawnwood exports to the UK. The result does not agree with the earlier studies, which indicate that the trade flows of forest products are not very sensitive to the exchange rate variations (Buongiorno et al. 1988 and Uusivuori and Buongiorno 1990). The previous results are however scanty and they concern exports to the US market.

The responsiveness of trade flows to exchange rate changes is connected with the magnitude of PT. For example, a low PT indicates that trade flows remain relatively insensitive to exchange rate changes (Menon 1995). Thus, the above-mentioned results of Buongiorno et al. (1988) and Uusivuori and Buongiorno (1990) indicate low PT for exports to the USA. Several studies have estimated low PTs (e.g. Yang 1997) for U.S. imports of many commodities. This indicates that trade flows directed to US market have not been very sensitive to exchange rates.

The pass-through coefficient obtained from the restricted price equation (16) was close to unity (0.92) and it could be further restricted to one. This means that exchange rate pass through in

Finnish sawnwood price expressed in pounds sterling has been large. The result is consistent with the relatively large exchange rate elasticity obtained from the demand equation. However, the lack of proper data on sawnwood production costs may have compromised the measurement of PT.

Several earlier findings for forest industry products indicate relatively large PTs with the exception of US imports. For example, Uusivuori and Buongiorno (1991) estimated PT to be between 0.79 and unity for US lumber exports to Japan. According to Menon (1993a), PT was 0.80 for Australian imports of wood products and 0.45 for paper products. The results of Vesala (1992) showed that PT for the export prices of Finnish paper manufacturers varied between 0.66 and 0.69 in western Europe and between 0.16 and 0.30 in the USA.

A large PT for Finnish sawnwood exports means that variations in exchange rates are almost completely passed through to the foreign currency prices in the long-run and the markup is kept unchanged. For example, depreciations and devaluations of the FIM have lowered the relative Finnish price and improved competitiveness and market share in the UK. Because prices expressed in foreign currency change as a result of an exchange rate change, a large PT also implies imperfect competition. Also in this respect, the results are consistent with earlier results implying imperfect competition and rejection of the law of one price for UK sawnwood imports (Hänninen 1998).

The resulting large PT indicates that Finnish sawnwood exporters have made use of devaluations to increase their market shares and export quantities but not necessarily their profit margins. If EMU is realized and Finland joins, the possibility of using exchange rate policy to improve Finnish competitiveness and market share will disappear. Even outside the EMU, the scope for national exchange rate policy would be reduced (Hetemäki et al. 1997). Hence, other means must be found to adjust to future disturbances caused by demand and price decreases in the world market.

One way to improve Finnish competitiveness during recessions could be to adjust production costs. However, a rapid adjustment is problem-

atic due to the relatively long-term price contracts that exist in the Finnish input markets. Hence, it is also necessary to differentiate Finnish products, for example via customer orientation, established customer relationships and by product planning. Because PT seems to be higher for wood industry products than for paper products, according to the earlier results, the sawwood industry may have more difficulty than the paper industry in adjusting to the EMU environment.

For further research, the short-run dynamics of exchange rate changes vis-à-vis Finnish sawwood prices would provide a useful and interesting challenge. Earlier studies of pass-through indicate that price adjustment does not necessarily happen immediately (Dornbusch 1987). The pass through effects could also be examined with respect to other forest products and other currencies, for example the Swedish krona, which has played an important role in western European sawwood trade, particularly in the 1990s.

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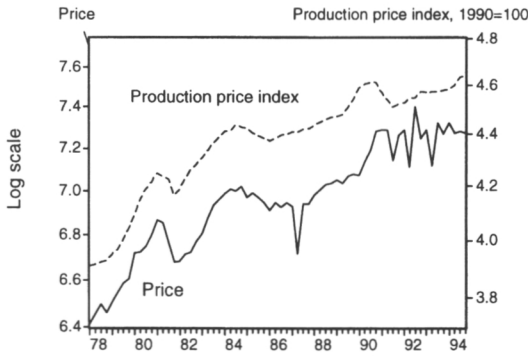
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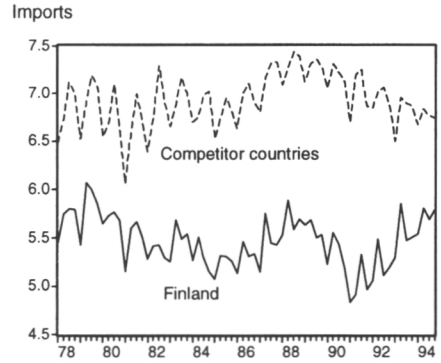
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Appendix 1. Quarterly time series data in logarithms covering the period 1978–1994.

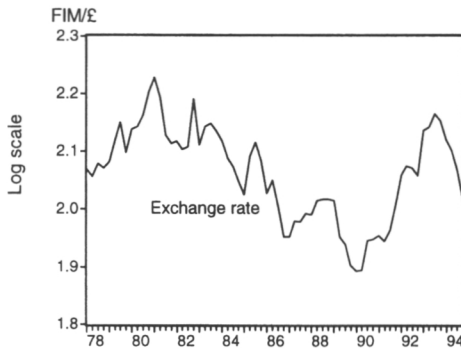
1A. Price of Finnish sawwood in the UK (FIM/m³) and Finnish production price index (1990=100).



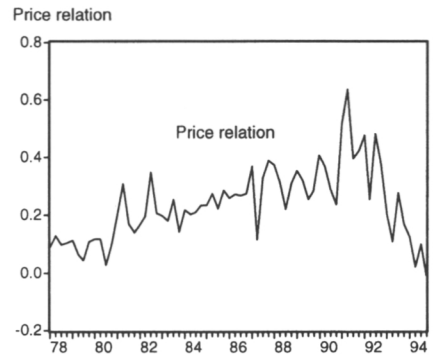
1B. UK sawwood imports from Finland and competitor countries, 1000m³



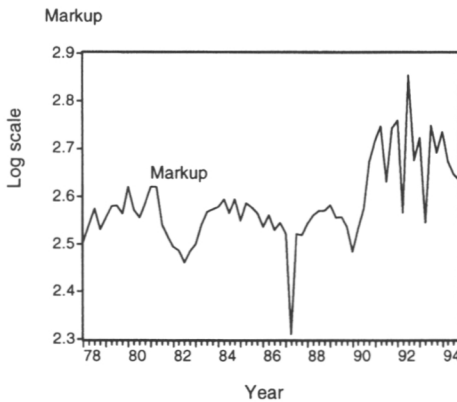
1C. Exchange rate, FIM/£.



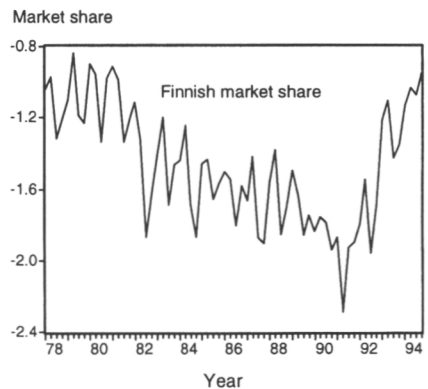
1D. Price relation between Finnish price (£/m³) and the competitors' price (£/m³).



1E. Markup. Relation between Finnish price (FIM/m³) and production price index (1990=100).



1F. Finnish market share of the UK sawwood imports.



V

LONG-RUN PRICE EFFECTS OF EXCHANGE RATE CHANGES IN FINNISH PULP AND PAPER EXPORTS

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ABSTRACT

This paper presents long-run exchange rate elasticities of Finnish newsprint and pulp export prices in the United Kingdom (UK) and Germany over the period 1980-1994. The elasticities were estimated for a markup pricing model using Johansen's cointegration method. The magnitudes of the elasticity estimates obtained by imposing price homogeneity indicate that pass-through of exchange rate changes to export prices has been incomplete in both markets. For newsprint, the pass-through elasticity was 0.60 in the UK and 0.46 in Germany, which means that about one half of the changes in the exchange rates is reflected in prices as expressed in the destination countries' currencies and about a half in prices measured in Finnish markka. Instead, in pulp exports the pass-through elasticities differed from each other between the two destination countries more clearly than in newsprint exports indicating a different pricing behaviour of Finnish pulp exporters in the UK and Germany when exchange rate changes. Pass-through was small in the UK pulp market (0.07) as compared to Germany (0.68).

Keywords: Exchange rate pass-through, international trade, newsprint, wood pulp, Johansen's cointegration method, United Kingdom, Germany

I. INTRODUCTION

A distinctive feature in the economies of the Nordic countries and Canada is the large share of forest industry in the total manufacturing and exports of goods. In Finland, forest industry products account for about one third of the total export earnings. Cyclical fluctuations of the forest industry are detectable in the overall performance of the Finnish economy. Finnish currency (markka) has been devaluated several times, which may have improved the profitability and competitiveness of Finnish export industry in relation to its main competitors, e.g., Sweden and Canada. However, empirical evidence on the effects of devaluations on the exports of forest industries is largely missing. Knowledge on the exchange rate effects is important in analysing possible implications of Finland's participation in the third stage of the European Monetary Union (EMU). Adjustment difficulties to the new market environment may arise for Finnish pulp and paper trade in the euroarea, if the exchange rate has had an important effect on Finnish export pricing and exports. Examining the effects of exchange rate changes on product prices provides also information on producers' pricing behaviour in the European pulp and paper markets.

Exchange rate effects are examined in the present study using the concept of pass-through (PT), which can be defined as the extent to which exchange rate changes are reflected in the export price. PT is said to be complete ($PT=1$) when the exporter does not adjust prices in his home currency, and exchange rate changes are fully reflected in foreign currency-expressed export prices. By contrast, there is no PT ($PT=0$), if the home currency-expressed export prices fully adjust to exchange rate fluctuations. PT may also be incomplete (partial), which means that exchange rate changes are partly reflected in foreign currency prices and partly in the exporters' home currency prices ($0 < PT < 1$).

PT also produces information on the degree of price competition in the markets (e.g., Athukolara 1991). In the earlier literature, incomplete PT is commonly attributed to imperfect competition and the use of market power. Thus, the magnitude of PT indicates the effects of depreciation/appreciation of the markka on Finland's price competitiveness in forest products' markets.

The interest to study PT originates from the contemporary debate on the efficiency of exchange rate policy in balance of payments adjustment. In most previous studies PT is found to be incomplete (see, e.g., the survey by Menon 1995). The rate of PT is also found

to differ between countries indicating different pricing behaviour. Some studies indicate that PT for total export and import prices tends to be higher for smaller more-open economies than for larger less-open economies (Japan, USA) (e.g., Kreinin 1977, Spitaeller 1980). However, results from the multicountry studies indicating the PT based on the size and openness of the country are conflicting (see e.g., Menon 1995, Khosla and Teranishi 1989).

The estimates for PT have also been found to vary widely by the industry (e.g., Feenstra *et al.* 1996), which means that it is necessary to study the concept at the commodity level. Most studies concern aggregate exports or imports, while more recent research analyse PT also at the industry-level (see, e.g., Hooper and Mann 1989, Menon 1995). The response of export price to changes in exchange rates depends on several factors that affect through the impact that exchange rates have on marginal costs and markups of price over marginal cost. The adjustment of markups in response to exchange rate change is referred to as pricing to market (Krugman 1987). Studies that examine pricing to market (PTM) find that the degree of PT varies across export markets. They suggest that firms price discriminate across export markets by varying the degree of exchange rate PT (e.g., Athukolara and Menon 1995). PTM approach has been applied by e.g., Knetter (1989, 1993), and Pick and Park (1991) in estimating pass-through elasticities.

Studies concerning PT of forest products are few. However, the opinion that changes in exchange rates play a key role in international trade in forest products is widely held, and a devaluation is expected to increase exports of forest products (e.g. McCarl and Haynes 1985). From the previous studies can be mentioned Vesala (1992), who studied exchange rate PT for Finnish paper exports. PT was found to be higher for exports to Europe (in the range 0.66-0.69) than for exports to the USA (in the range 0.16-0.30), which may be due to the larger domestic paper production in the USA than in Europe. Uusivuori and Buongiorno (1991) found incomplete PT in forest products exports from the USA to Europe and Japan for most of the product categories. According to Alavapati *et al.* (1997) exchange rate PT was small in Canadian pulp export price indicating that Canadian exporters do not have significant market power in pulp exports. The PT estimate of Menon (1993) for Australian imports of paper was 0.45. The results of the previous studies indicate a larger PT for wood industry products than for paper industry products (Uusivuori and Buongiorno 1991, Menon 1993 and Hänninen 1998).

Previous models have often been bivariate, where a price variable has been regressed on the exchange rate using ordinary least squares (OLS) method (e.g., Kreinin 1977, Knetter 1989, 1993, Pick and Park 1991, Spitaeller 1980, Uusivuori and Buongiorno 1991). Using OLS in the estimation of PT has been criticised (e.g., Menon 1995), because the data are usually nonstationary (trended) and thus the OLS estimates may suffer from “spurious regression” problems. Multivariate approaches in estimating PT have commonly applied a markup pricing equation (see e.g., Dornbusch 1987, Hooper and Mann 1989, Athukolara and Menon 1994 and 1995, Hung *et al.* 1993, Kuismanen 1995, Naug and Nymoen 1996, Vesala 1992, Hänninen 1998). The models of these studies are based on the assumption that the markup is variable and affected, for example, by the foreign competitor's price and foreign demand pressure. In the study of Rickerbie (1992) estimable equation for analysing PT was derived from the reduced form of an export demand and supply model.

The analysis of the present study is based on a markup model for Finnish exports of the two most homogenous paper industry products, i.e. newsprint and pulp to the UK and Germany. The Johansen (Johansen 1988, Johansen and Juselius 1990, 1992) multivariate cointegration method for nonstationary time series data is applied in the estimation. This method has not yet been widely used in PT studies (see, however, Menon 1993, Naug and Nymoen 1996, Alavapati *et al.* 1997, Hänninen 1998), although it has many advantages compared to the traditional cointegration methods (e.g., Engle and Granger 1987). For example, in multivariate models it is possible to test the number of cointegration vectors among the variables and to test the coefficient restrictions in the same cointegration framework.

II. THEORETICAL MODEL

A model for estimating PT for Finnish paper and pulp exports to the UK and Germany is a markup model that is previously used in analysing PT e.g., by Hooper and Mann (1989), Athukolara (1991) and Hung *et al.* (1993). In deriving the price equation, it is assumed that a representative Finnish exporter produces exclusively for imperfectly competitive UK or German markets, employs a constant returns to scale technology and unit production cost, *C_f*. The representative Finnish firm maximizes profit by taking the competitors' price and

supply of newsprint (or pulp) as given and by setting the price in Finnish markkaa (P_f) as a constant markup over unit production costs (C_f). With X_f denoting exports, the exporter's profit is V_f defined as

$$(1) \quad V_f = (P_f - C_f) X_f$$

Profit maximization yields

$$(2) \quad P_f = C_f \eta / (\eta - 1)$$

where η is the price elasticity of demand. According to Hung *et al.* (1993), a more general case, in which the competitor's price affects the exporter's price, can be presented using the concept of a variable markup, defined by assuming that the coefficient η depends on, among other things, price competitiveness in the export market. The competitiveness can be described as the relative price $(P_o ER)/P_f$ where $(P_o ER)$ is the competitor's price in the exporter's currency and ER is the exchange rate (FIM/GBP in the UK and FIM/DEM in Germany). Thus the price elasticity of demand is

$$(3) \quad \eta = \eta((P_o ER)/P_f)$$

From (2) and (3) the pricing behavior of the profit maximizing exporter can be described as a variable markup over the unit cost:

$$(4) \quad P_f = \phi C_f$$

The variable markup, ϕ , depends on the relative price:

$$(5) \quad \phi = \phi((P_o ER)/P_f) = \phi'((P_o ER)/P_f)^\theta$$

where θ is the relative price elasticity of markup and $\theta \geq 0$. The constant markup is obtained if $\theta = 0$ and $\phi' = \eta/(\eta - 1)$. The second equality in (5) derives from the log-linear approximation of the nonlinear function ϕ . Substituting (5) into (4) and taking a logarithmic

transformation, a relation for the price of Finnish newsprint (or pulp) in the export market is obtained:

$$(6) \quad pf = \delta + (1-\gamma)(er + p_o) + \gamma cf + u$$

where $\gamma = 1/(1+\theta)$, δ is a constant and u is the disturbance term that captures all other possible factors. Symbol γ ($0 < \gamma \leq 1$) indicates the PT coefficient ($PT = -(\partial(pf - er)/\partial er) = \gamma$). The other symbols are the same as above. Lower-case variables denote logs of corresponding upper-case variables. The export price, pf , should be homogenous of degree zero in the exchange rate and the competitor's price, and the equality restriction is imposed on the coefficients of p_o and er in the estimation.

If the coefficient of Finnish cost, γ , is unity in (6), Finnish export price in markkaa (FIM) is determined solely by Finnish production cost. Thus, the markup is held constant after a change in the exchange rate. In this case, the changes in er (FIM/GBP, FIM/DEM), will completely pass through on the Finnish export price denoted in foreign currency and $PT=1$. If $\gamma = 0$, the changes in er (FIM/GBP, FIM/DEM) are absorbed by the variable markup and $PT=0$. Thus, export prices in foreign currency do not change (net of any effect exchange rate changes may have on prices through variations in input costs). This indicates perfect competition in the market. Between the two extremes, PT may be incomplete ($0 < PT < 1$). In this case, the changes in er will change the export price, both in FIM and in foreign currency and exchange rate changes are partially absorbed into the variable markup.¹

III. DATA

The data of the study are quarterly, seasonally unadjusted and cover the period from the first quarter of 1980 to the fourth quarter of 1994 ($N=60$). The Finnish export prices (pf)

¹ The PT may also be affected by factors other than these included in the model (6). If Finnish production costs include an important imported component, it is evident that the degree of PT of exchange rate changes is less than one, even if $\gamma=1$. In the Finnish pulp and paper industry, the inputs are for the most part of domestic origin, which indicates that this effect is not very important.

are described by the average FOB (free on board) export unit values of newsprint (CN 48.01) and pulp (CN 470321 and CN 470329) in Finnish markkaa (FIM) and they are obtained from the customs statistics of Finland (The Board of Customs). The foreign competitor's prices (p_o) are the average CIF (including cost, insurance and freight) unit values (GBP/metric tn and DEM/tn) of total newsprint and pulp imports (excluding imports from Finland) to the UK and Germany.² The source of this data are the customs statistics of the UK (CSO) and Germany (Statistisches Bundesamt). Exchange rates (er) between the Finnish markka and GBP and DEM are quarterly averages and they were obtained from the Bank of Finland. The producer price index (1990=100, for commodity group SIC 341) obtained from Statistics Finland is used to describe Finnish production costs (cf) for the pulp and newsprint industry, because better cost data are not available. This may cause uncertainty in estimating PT, which must be taken into account in drawing conclusions.

The UK and Germany are important markets for Finnish newsprint and pulp, accounting for about 40 % of newsprint and 60 % of pulp exports from Finland. Finnish average market share from total newsprint imports has been about 28 % in the UK and 22 % in Germany during the period studied. Finland has lost market share in both countries to especially Sweden. This development has been offset by the acquisition of production capacity owned by Finnish forest industry companies in the UK market during the 1980s and in Germany during the 1990s. In pulp imports to the UK and Germany Finnish average market shares have been around 15 %.

In Figure 1A, the export prices of Finnish pulp and newsprint denoted in the importers' currency in the UK and Germany are graphed together with the exchange rates. It can be seen that prices (in GBP and DEM) and exchange rates fluctuate against each other, which suggests that there has very likely been some degree of exchange rate PT in both markets; markka depreciations have lowered the Finnish price and markka appreciations have raised it.

The long-run development of the exchange rates FIM/GBP and FIM/DEM differed from each other in the 1980s. The Finnish markka appreciated with respect to GBP and

² The product codes of newsprint are for British imports SITC 641.1 and for German imports CN 4801 0010 including item CN 4801 650 for year 1980, item CN 4801 790 for years 1981–87 and item CN 4801 0090 for 1988 onwards. Product codes of pulp (bleached sulphate pulp) are for British imports SITC 251.51 and SITC 251.52 and for German imports CN 470321 and CN 470329.

depreciated with respect to DEM. In the early 1990s the exchange rates have behaved similarly and the large devaluations of the markka in 1991 and 1992 can be seen in the graphs. The markup over production costs ($pf - cf$) in Figure 1B indicates that the profitability margin has fluctuated over time. All in all, the development of the variables of the study indicates that studying the degree of PT is important in modelling the international trade of pulp and paper.

IV. ESTIMATION METHOD AND RESULTS

The present study uses Johansen's cointegration method, which is suitable for estimation purposes when the variables are nonstationary. Compared to the traditional cointegration methods (e.g., Engle and Granger 1987) Johansen's method has several advantages. Unit roots are explicitly incorporated in the specification, and the method takes into account short-run dynamics in estimating the cointegration vector. Additionally, linear restrictions of the coefficients, e.g., price homogeneity, can be tested in the same cointegration framework. The Johansen's method is briefly summarized below, but a comprehensive description of estimating cointegration vectors and testing hypotheses can be found in Johansen (1988) and (1995), see also Johansen and Juselius (1990, 1992). The method uses a statistical model that is a p -dimensional VAR process of order k , where p is the number of variables. The VAR model can be reparameterized in error correction (EC) form as:

$$(7) \quad \Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu + \Phi D_t + \varepsilon_t, \quad t=1, \dots, T,$$

where Δx_t is a $I(0)$ vector of the first-differenced variables, μ is a vector of constants, D_t is a seasonal dummy and k is the lag length ($k = 1, \dots, N$). In (7) constant μ can be restricted to the cointegration space, if no linear trend is present in the data series. Π is the matrix of long-run coefficients of the lagged levels. It can be decomposed into a matrix of loadings, α , and a matrix of cointegrating vectors, β , i.e., $\Pi = \alpha\beta'$. The number of cointegration vectors that exist among the variables is determined by estimating the rank of the matrix Π . Johansen (1988) formulates two likelihood ratio tests for the cointegration rank, a

maximum eigenvalue test and a trace test, from which the trace test results are reported here.

Provided that the data series are cointegrated, it may be possible to estimate a cointegration vector the coefficients of which describe the long-run equilibrium relationship implied by equation (6) for each destination country and product. The estimations are based on the VAR model given by equation (7) with four individual equations. The suitable lag length, k , of the VAR model was determined by diagnostic testing of the data in each case. In addition, the Schwarz (SC) and Hannan-Quinn (HQ) information criteria (using likelihood ratio tests for sequential decreases in the number of lags from 5 to 1 in each of the systems) were applied in the determination of the lag length of each of the VAR models (see Doornik and Hendry 1994).

When the number of cointegration relations, r , is determined it is possible to test hypotheses on the long-run matrix, $\Pi = \alpha\beta'$. The present study tests restrictions on the coefficients β_{ij} of the cointegration vectors. Homogeneity between Finnish export price, pf , and the other nominal variables, was tested by restricting the respective coefficients accordingly. The markup relationship was also tested by examining if the estimate for γ (see equation 6) could be restricted as unity in the cointegration vector.

Before proceeding with cointegration estimation, the individual time series were tested for nonstationarity using Augmented Dickey Fuller tests (Dickey and Fuller 1979).³ Test results in Table 1 indicate that all the time series are nonstationary and integrated of order one. The cointegration approach is thus suitable for modelling the long-run exchange rate PT in British and German markets of pulp and paper.⁴

³ Also Phillips-Perron (Phillips and Perron 1988) modification of the unit root tests were performed. As the results were found to be uniform regarding the null hypothesis of nonstationarity, the results are not presented here (available from authors upon request).

⁴ The VAR-models were tested for stability by using Chow-tests and examining the recursive residuals of the four models systems. According to the resulting recursive constancy statistics, there were no problems with the stability of the models. Regarding the cointegrating relationships, also their stability was examined with plotting recursive eigenvalues of individual long run vectors. No signs of instability was detected here either.

Results for Johansen's cointegration rank test for newsprint price model of the UK market are presented in Table 2. The VAR model with seasonal dummies, restricted constant and two lags was found to be sufficient to remove residual autocorrelation of individual equations (diagnostic test results are presented in Appendix 1A). The cointegration rank was found to be one using the trace test with and without the degrees of freedom correction.

Also in the German data, a VAR model with seasonal dummies, restricted constant and two lags was found to be a sufficient description of the data generating process (see Appendix 1A). According to the cointegration rank test (Table 2), the test value for $r = 0$ is only slightly lower than the critical value, so we can reject $r = 0$ and accept $r = 1$.

Economic theory requires homogeneity of prices following equation (6). The homogeneity in nominal variables pf , cf and $(p_o + er)$ implies that the coefficients of p_o and er should be equal and that the coefficient of cf should equal the difference between the coefficients of pf and er in the cointegration vector. This assumption was tested by restricting the coefficients of the UK and German cointegration vectors accordingly. Long-run price homogeneity could be accepted (Table 3) and the restricted cointegrating vectors can be written as

$$(8) \quad \text{UK:} \quad \beta' : 1.00pf = +0.40(er+p_o) + 0.60cf + 1.85,$$

$$(9) \quad \text{Germany:} \quad \beta' : 1.00pf = +0.54(er+p_o) + 0.46cf + 1.32,$$

where pf is Finnish price (FIM/tn), er is exchange rate, cf is Finnish production cost and p_o is competitors' price (GBP/tn, DEM/tn). The relation (8) indicates that from a 10 percent change of the exchange rate, 6 percent would in the long run be reflected in the Finnish export price in pounds sterling and 4 percent in the markup over domestic costs. If the exchange rate changes were fully reflected in GBP prices, PT would be complete and prices in FIM would remain unchanged. This hypothesis was also tested, but the restriction in the long-run newsprint price relation could not be accepted in the UK (Table 3). For Germany the relationship (9) indicates that in the long-run equilibrium less than one half of an exchange rate change is reflected in DEM and over a half in FIM prices. Consequently, the

markup assumption, i.e. the restriction of production cost coefficient (γ from equation 6) to one (Table 3) could not be accepted for Germany either.

Exchange rate effects in Finnish pulp prices

For the pulp price model of the UK, the VAR model with seasonal dummies, restricted constant and four lags was found to be a sufficient description of data generating process (see Appendix 1B). Johansen's cointegration rank test in Table 4 indicate the cointegrating rank to be one (at 1 % level) or two (at 5 % level). However, since the degrees of freedom corrected trace test (not reported here) indicated only one cointegrating vector, we proceed according to the theoretical model and identify the pass-through coefficient through imposing price homogeneity in the UK cointegration vector (Table 5).

Also in the German data, a VAR model with four lags, seasonal dummies and restricted constant was found to be a sufficient description of the data generating process (see Appendix 1B). According to the cointegration rank test (Table 4), the trace test result for $r \leq 2$ is above the critical value at 5 % level. However, when using the degrees of freedom corrected version of this test (not reported), the result $r=1$ is obtained. We proceed accordingly and make the coefficient restrictions under rank $r=1$ also on the German cointegrating vector. The markup pricing relations in both countries can be written as

$$(10) \text{ UK: } \beta' : 1.00pf = +0.93(er+p_0) + 0.07cf + 0.18$$

$$(11) \text{ Germany: } \beta' : 1.00pf = +0.32(er+p_0) + 0.68cf + 2.25,$$

where the symbols are the same as above. In the UK cointegrating vector the homogeneity restriction was accepted and the coefficient for long run exchange rate pass-through of 0.07 is obtained for the pulp price (Table 5). Even the zero restriction for the PT coefficient could be imposed for pulp price in the UK, indicating that the Finnish exports of pulp have been insensitive to exchange rate changes. Instead, the restriction of markup relation to express complete pass through (i.e. $PT=1$) is rejected in the UK data.

For Germany, the likelihood ratio test statistic for the restricted price relationship was rather close to 5 % critical value, so we imposed the restriction. According to (11) close to

70 % of an exchange rate change in the German market was thus reflected in DEM prices and over 30 % in prices denoted in FIM in the long-run equilibrium. Consequently, the markup assumption, i.e. the restriction of production cost coefficient to one (Table 5) could not be accepted for German pulp price. If the magnitude of PT coefficient is interpreted in terms of market competition, the low (even zero) value of it reflects that the UK pulp market is more competitive than German market.

V. CONCLUDING REMARKS

This paper examines the long-run exchange rate pass-through (PT) of Finnish pulp and newsprint export prices in the UK and Germany. According to the estimation results, pass-through effects have been incomplete for both product groups and markets. Thus, changes in the exchange rates are partly reflected in Finnish pulp and paper export prices in foreign currency and partly in the markup over production costs. For newsprint, the elasticity estimates were relatively close to each other in both destination countries (in the range 0.46–0.60), while for pulp, the UK elasticity (0.07) was small as compared to Germany (0.68).

The estimated PT coefficients indicate that devaluations have boosted Finnish export demand of paper in the UK and German markets, and that the competition between Finland and the other exporters in these markets has been imperfect. Moreover, Finnish paper exporters appear to have pursued a midway pricing strategy aimed at maintaining market shares and profitability as the exchange rate changes (see also Athukorala and Menon 1995). In contrast, the pricing of Finnish pulp exports has been more clearly destination specific as compared to paper.

The results of this study agree with Vesala (1992), whose PT estimate for Finnish paper exports to Europe was 0.66. For newsprint, the present study estimated a slightly lower PT coefficient in Germany than in the UK, which may be related to the smaller market share in Germany than in the UK. Germany is also a large newsprint producer, whose domestic producers may have market power in relation to the foreign importers. One of the underlying reasons for incomplete PT in the UK and Germany may be the location of Finnish paper production capacity in these countries. According to Gron and Swenson

(1996), firms' export prices are unlikely to change one-for-one with exchange rate movements if they have production capacity across borders. This may also have affected the magnitudes of the PT estimates for pulp, which seem not be connected with Finnish market shares in the same way as those of the newsprint. Other things related to, for example, the use of pulp as a rawmaterial of paper and a possibility to transfer pricing of pulp in the paper mills owned by Finnish forest industry in the customer countries may have effects on the PT estimates of pulp.

The implication of this study for Finland's participation in stage three of EMU is that some adjustment difficulties could arise in pulp and paper industry, because the exchange rate between the Finnish markka and the main importing countries' currencies have affected exports. Along with euro, Finland and Germany will have a common currency, while with respect to the UK the exchange rate will be determined between euro and GBP. Basing on the PT elasticities the effects of EMU membership on the competition for market share are, however, likely to be smaller in Finnish pulp and paper industry than in Finnish sawnwood industry. When Finland joins the EMU, national exchange rate policy can no longer be used to improve Finnish price competitiveness. Hence, other means must be found to adjust to future disturbances caused by shocks in demand and world market prices.

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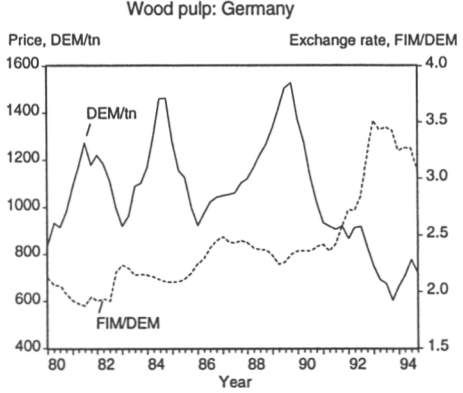
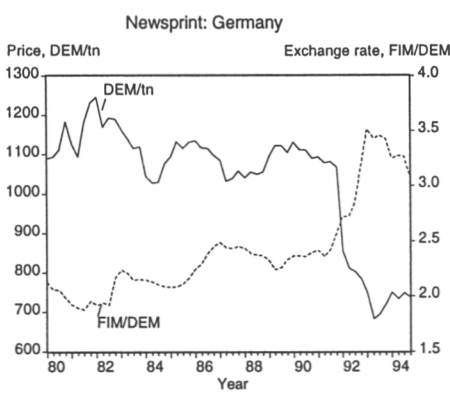
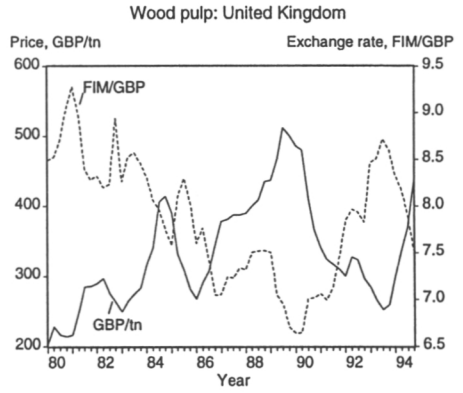
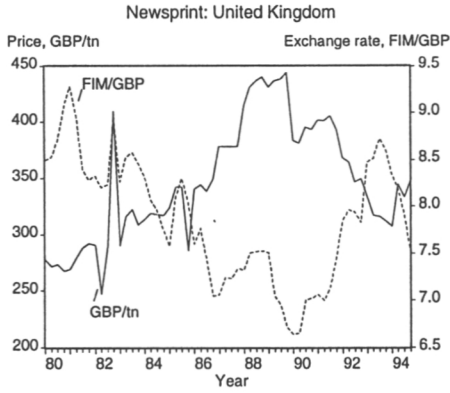


Figure. 1A. Nominal CIF - prices of Finnish newsprint and wood pulp in the UK (GBP/tn) and Germany (DEM/tn) with the exchange rates, FIM/GBP and FIM/DEM.

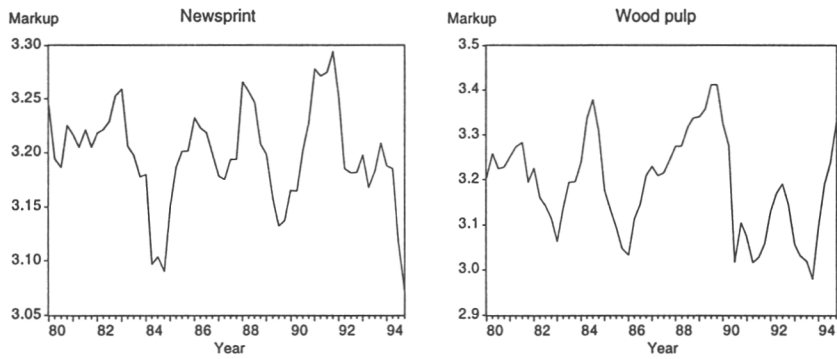


Figure. 1B. Markup. Relation between Finnish unit price of exports (FIM/tn) for newsprint and pulp and production price index (1990=100) for paper and pulp industry.

Table 1. Results for unit root tests.

	Level ¹	First difference
Newsprint: export price from Finland	-1.78 (T, N=0)	-6.43**
Wood pulp: export price from Finland	-1.65 (T, N=0)	-4.91**
Production cost	-2.62 (N=0)	-4.47**
Exchange rate, FIM/GBP	-2.34 (N=0)	-6.57**
Exchange rate, FIM/DEM	-1.61 (T, N=0)	-4.35**
Newsprint: competitor's price in UK	-2.56 (N=0)	-8.62**
Newsprint: competitor's price in Germany	-0.93 (N=1)	-6.16**
Wood pulp: competitor's price in UK	-3.16 (T, N=0)	-8.33**
Wood pulp: competitor's price in Germany	-1.53 (T, N=0)	-3.98**

¹ For critical values, see Dickey and Fuller (1979).

Note: Constant or constant and trend (T) included in test equation and N= number of lags needed in test equation to remove autocorrelation. *(**) denotes rejection of null hypothesis of nonstationarity at 5 % (1 %) level.

Table 2. Results for the cointegration rank test in British and German newsprint models.

$H_0: r \leq i$	United Kingdom		Germany		95 % critical value
	Eigenvalues λ_i	Trace test statistic	Eigenvalues λ_i	Trace test statistic	
$r=0$	0.43	59.8	0.39	52.5	53.1
$r \leq 1$	0.31	31.5	0.23	24.5	34.9
$r \leq 2$	0.17	13.2	0.15	10.2	20
$r \leq 3$	0.08	4.0	0.02	1.03	9.2

Table 3. Tests for restrictions on the newsprint price relations in the UK and Germany.

	p_f	er	c_f	p_o	Constant	$\chi^2(2/3)^*$
UK:						
Price	1.00	-0.40	-0.60	-0.40	-1.85	1.41
homogeneity						[0.49]
Complete	1.00	0	-1.00	0		18.8
pass-through						[0.000]
Germany:						
Price	1.00	-0.54	-0.46	-0.54	-1.32	5.25
homogeneity						[0.07]
Complete	1.00	0	-1.00	0		9.14
pass-through						[0.03]

* χ^2 critical value for at 5 % level is 5.99 with two degrees of freedom and 7.82 with 3 d.g.f.
 Symbols: p_f denotes the export price from Finland, er is the exchange rate, c_f is Finnish production cost and p_o is the competitor's price.

Table 4. Results for the cointegration rank test in British and German pulp models.

$H_0: r \leq i$	United Kingdom		Germany		95 % critical value
	Eigenvalues λ_i	Trace test statistic	Eigenvalues λ_i	Trace test statistic	
$r=0$	0.42	71.6	0.45	80.0	53.1
$r \leq 1$	0.34	40.7	0.36	46.9	34.9
$r \leq 2$	0.19	17.7	0.24	21.8	20.0
$r \leq 3$	0.10	5.9	0.11	6.6	9.2

Table 5. Tests for restrictions on the pulp price relations in the UK and Germany.

	p_f	er	c_f	p_o	Constant	$\chi^2(2/3)^*$
UK:						
Price	1.00	-0.93	-0.07	-0.93	-0.18	0.06
homogeneity						[0.97]
Complete	1.00	0	-1.00	0		19.93
pass-through						[0.000]
Germany:						
Price	1.00	-0.32	-0.68	-0.32	-2.25	6.74
homogeneity						[0.03]
Complete	1.00	0	-1.00	0		17.15
pass-through						[0.000]

* χ^2 critical value for at 5 % level is 5.99 with two degrees of freedom and 7.82 with 3 d.g.f.

Note: For symbols, see Table 3.

APPENDIX 1A.

Misspecification tests for residuals from Johansen's cointegration estimation of British and German newsprint models with two lags, seasonal dummies and restricted constant.

Equation	Tests for residuals and standard errors			
	Autocorrelation ^{a)}	Heteroskedasticity ^{b)}	Normality ^{c)}	Standard errors
	$F_{AR(4,40)}$	$F_{ARCH(4,36)}$	$\chi^2(2)$	σ_e
UK:				
Δp_f	0.16 [0.96]	0.58 [0.68]	1.84 [0.40]	0.02
Δer	0.86 [0.50]	0.16 [0.96]	7.36 [0.03]	0.03
Δc_f	1.03 [0.40]	0.07 [0.99]	3.23 [0.20]	0.02
Δp_o	1.30 [0.29]	0.05 [0.99]	47.19 [0.00]	0.01
System:	$VF_{AR}(64,100)=1.17[0.24]$		$V\chi^2(8)=39.26 [0.00]^*$	
Germany:				
Δp_f	0.29 [0.89]	0.64 [0.64]	0.84 [0.65]	0.02
Δer	0.73 [0.58]	2.19 [0.09]	5.23 [0.07]	0.03
Δc_f	2.53 [0.06]	0.49 [0.74]	3.66 [0.16]	0.02
Δp_o	1.77 [0.15]	5.48 [0.00]^*	5.77 [0.06]	0.02
System:	$VF_{AR}(64,100)=1.25[0.16]$		$V\chi^2(8)=11.79 [0.17]$	

Note: Values in square brackets are marginal significance levels and * indicates that the null hypothesis is rejected at the 5 percent level. ^{a)} Autocorrelation of the residuals of individual equations and a whole system was examined using the F-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged dependent variables. ^{b)} Heteroskedasticity was tested using the F-form of the LM test against 4th order autoregressive conditional heteroskedasticity. ^{c)} Normality of the residuals of individual equations and a whole system was tested with the Doornik-Hansen test (Doornik and Hendry 1994). For further detail and test references, see Doornik and Hendry (1994).

Symbols: p_f denotes the export price from Finland, er is the exchange rate, c_f is Finnish production cost and p_o is the competitor's price.

APPENDIX 1B.

Misspecification tests for residuals from Johansen's cointegration estimation of British and German pulp models with four lags, seasonal dummies and restricted constant.

Equation	Tests for residuals and standard errors			
	Autocorrelation ^{a)}	Heteroskedasticity ^{b)}	Normality ^{c)}	Standard errors
	$F_{AR}(4,32)$	$F_{ARCH}(4,28)$	$\chi^2(2)$	σ_e
UK:				
Δp_f	0.07 [0.99]	0.88 [0.48]	9.06 [0.01]*	0.07
Δer	3.25 [0.02]*	0.21 [0.93]	3.69 [0.16]	0.03
Δc_f	1.60 [0.19]	0.38 [0.82]	3.49 [0.17]	0.01
Δp_o	1.38 [0.26]	3.71 [0.02]	39.4 [0.00]*	0.13
System:	$VF_{AR}(64,68) = 0.97 [0.55]$		$V\chi^2(8)=69.34 [0.00]*$	
Germany:				
Δp_f	0.13 [0.97]	1.30 [0.29]	3.97 [0.14]	0.06
Δer	0.69 [0.60]	1.15 [0.36]	5.83 [0.05]	0.03
Δc_f	1.64 [0.18]	0.26 [0.90]	2.73 [0.26]	0.01
Δp_o	2.55 [0.06]	0.35 [0.84]	1.02 [0.60]	0.04
System:	$VF_{AR}(64,68)=1.07 [0.39]$		$V\chi^2(8)=22.78 [0.01]*$	

Notes: Values in square brackets are marginal significance levels and * indicates that the null hypothesis is rejected at the 5 percent level. ^{a)} Autocorrelation of the residuals of individual equations and a whole system was examined using the F-form of the Lagrange-Multiplier (LM) test, which is valid for systems with lagged dependent variables. ^{b)} Heteroskedasticity was tested using the F-form of the LM test against 4th order autoregressive conditional heteroskedasticity. ^{c)} Normality of the residuals of individual equations and a whole system was tested with the Doornik-Hansen test (Doornik and Hendry 1994). For further detail and test references, see Doornik and Hendry (1994).

For symbols, see Appendix 1A.



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