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MARKET IMPACTS AND POLICY IMPLICATIONS OF U.S.
RESTRICTIONS ON SOFTWOOD LUMBER IMPORTED
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MARKET IMPACTS AND POLICY IMPLICATIONS OF U.S. RESTRICTIONS ON SOFTWOOD LUMBER IMPORTED FROM CANADA

By

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ABSTRACT

MARKET IMPACTS AND POLICY IMPLICATIONS OF U.S. RESTRICTIONS ON SOFTWOOD LUMBER IMPORTED FROM CANADA

By

Jungho Baek

The softwood lumber trade dispute is the largest and most contentious dispute between the United States and Canada. The core issue of the dispute is Canada's subsidies to its lumber producers in the form of low stumpage fees and restrictions on log exports in British Columbia (BC). The U.S. government has imposed restrictions — the Memorandum of Understanding (MOU) during 1986:4-1991:3, the Softwood Lumber Agreement (SLA) during 1996:2-2001:1, and the current 27.2% tariff to regulated lumber flows from Canada.

With an expanded list of variables for the period of 1980:1-2001:4, this study developed a structural simultaneous system of equations (SSE) and a nonstructural vector auto-regression (VAR) model to analyze the U.S. lumber market structure and to assess the impacts of the U.S. trade restrictions.

The SSE and VAR models show that the MOU caused Canada's exports to decrease by 8-9% during 1986:4-1991:3, which in turn pushed up the domestic production by 3.0%, but it did not lead to higher lumber prices. However, the SLA did not have any significant impacts on Canadian exports, domestic production, and lumber prices during 1996:2-2001:1. In addition, the results show that, the price elasticities of domestic and export supply are smaller than those commonly used in evaluating the gains and losses of the U.S. restrictions, while such market factors as housing starts, log prices, and exchange rates have played major roles in shaping the lumber market.

Combining the modest market impacts of the restrictions and smaller price elasticities, therefore, leads to moderate welfare consequences of the trade restrictions. In addition, the simulated results under an export/import tax and a quota scheme show that from the U.S. perspective, a quota scheme is preferred to an export/import tax for restricting softwood lumber imported from Canada, as long as all Canadian provinces and all lumber products are included in the regime. In combination, these findings indicate the need not only for a broader and more dynamic perspective but also for a comprehensive approach to understanding the U.S.-Canada lumber trade and designing effective trade policies.

The assistance, support, and prayer of two
people made this dissertation possible.
I dedicate this work to them with all my love:
my wife, Eunji and my daughter, Jiyeon

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This study would not have been possible without the guidance and assistance of numerous individuals.

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Finally, I would like to thank my wife, Eunji, my daughter, Jiyeon (Jiya), and my parents, Donghak, Baek and Songja, Shin. Particularly, I would like to express my deep gratitude to my parents-in-law, Seunghan, Lee and Gunja, Song for their love, support, and prayers. I love you very much indeed.

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CHAPTER 1

INTRODUCTION

The softwood lumber trade dispute is the largest and most contentious dispute between the United States and Canada. The U.S. is the world's largest importer of softwood lumber, with more than 90% of its lumber imports coming from Canada. Canada is highly dependent on the U.S. lumber market, with more than 60% of its lumber production exported to the United States (Howard 2000). For decades the trade volume between the two countries has risen substantially, with Canada's share of the U.S. market having reached more than 30% in 2000, up significantly from about 15% in 1965. The core issues of the dispute are the U.S. claim that Canadian producers have expanded their U.S. market share with subsidized stumpage rates — fees paid by forest products companies to the provinces for the rights to cut trees — and restrictions on log exports in British Columbia.[1] As a result, over the last two decades the U.S. government has imposed trade restrictions to regulate lumber flows from Canada. The Memorandum of Understanding (MOU) required Canada to impose a 15% export tax on its lumber exported to the U.S. for the period between the fourth quarter of 1986 and the third quarter of 1991 (1986:4-1991:3). The Softwood Lumber Agreement (SLA) mandated Canada to charge at least $50 per thousand board feet (MBF) for its lumber exports in excess of 14.7 billion board feet (BBF) per year for the period between the second quarter of 1996 and the first quarter of 2001 (1996:2-2001:1). Currently, the U.S. government imposes tariffs of averaging 27.2% on imported Canadian lumber — 18.79% for countervailing duty (CVD) and 8.43% for anti-dumping duty (ADD), to protect the domestic lumber market.

1.1. Statement of Problem

Many scholars have studied the impacts of the trade restrictions on softwood lumber imported from Canada. Myneni et al. (1994) used a simultaneous system of equations (SSE) to estimate the welfare effect of the MOU (1986-1990). With the estimated price elasticities of demand (-0.10), supply (0.27), and Canadian export supply (0.19) they found that U.S. consumer lost $147.4 million/year (in 1982 dollars), whereas U.S. producer gained $109.5 million/year for the first 4 years under the MOU. As a result, the net U.S. loss was about $38.3 million/year. Other studies simply adopted a reduced-form equation (RFE) in their analyses. Wear and Lee (1993) used an RFE to investigate the market and welfare impacts of the MOU (1986-1990). They showed that the MOU reduced Canadian exports by 2.6 BBF/year — a 5% decline in Canadian market share, which in turn pushed up U.S. production by 1.83 BBF/year. Coupled with a price increase of $20/MBF annually, this resulted in a reduction in U.S. consumption of 0.76 BBF/year. As a result, U.S. consumers lost about $947.4 million/year (in 1982 dollars), whereas U.S. producers gained about $658.1 million/year, resulting in a net U.S. loss of $289.3 million/year for the first 4 years under the MOU.

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[1] British Columbia (BC) precludes exports of unprocessed logs to foreign countries. The U.S. producers insist that the restriction on log exports prevent non-BC producers from accessing logs harvested in BC. In addition, the log ban helps keep BC producers' wood costs artificially low (Ragosta and Clark 2000).

In contrast to the above moderate estimates, more recent RFE studies indicate substantial impacts from the U.S. trade restrictions. Lindsey et al. (2000) used the RFE for lumber prices to estimate the impacts of the MOU and the SLA. They found that the price effect of the trade restrictions was in the range of $50 to $80/MBF, which in turn pushed up the cost of new home prices by $800 to $1,300. Also, Zhang (2001) used an RFE to assess the price impact of the SLA. He found that under the SLA, the estimated annual change in lumber price was about $59.1/MBF, or 16%, on average, for the period 1996-2000. As a result, U.S. consumer lost about $12.5 billion (in 1997 dollars), whereas U.S. producers gained about $7.7 billion, resulting in a net U.S. loss of $4.8 billion for the first 4 years under the SLA.

In addition, some studies employed a vector auto-regression (VAR) model to investigate the impacts of macroeconomic variables on Canadian lumber exported to the U.S. Jennings et al. (1991) showed that macroeconomic variables, especially housing starts in North America have a significant effect on the Canadian lumber industry. Sarker (1996) showed that demand side factors, such as housing starts, disposable income, and wage rates have played key roles in affecting the lumber trade.

While previous studies have advanced our understanding of the market and welfare ramifications of the trade restrictions, they have some critical weaknesses. First, certain methodological weaknesses exist. Previous studies mostly employed an RFE to assess the impacts of the trade restrictions, which are vulnerable to problems of endogeneity, simultaneity, and collinearity (Wooldridge 2000). Endogeneity describes the presence of an endogenous explanatory variable, a variable that is correlated with the error term, either because of an omitted variable, measurement error, or simultaneity.

Simultaneity means that at least one explanatory variable in a multiple linear regression model is determined jointly with the dependent variable. Collinearity refers to the existence of correlation between explanatory variables. In addition, most of these studies employed small sample sizes, covering 20 to 30 annual observations (Wear and Lee 1993, Myneni et al. 1994, Lindsey et al. 2000). Coupled with the use of OLS estimation, these problems could cause coefficients of the RFE to be sensitive to its specification and even inconsistent (Wooldridge 2000).

Moreover, many studies have taken price elasticities from other sources for the welfare evaluations since the RFE model does not generate the relevant price elasticities. Considering the crucial role that the price elasticities play in generating the welfare gains and losses, it is not appropriate to take them from other sources. As noted by Boyd and Krutilla (1987), gains and losses to U.S. consumers and producers as well as to Canadian producers would be significantly altered by changes in the price elasticity of Canadian export supply alone. In fact, the price elasticity of Canadian export supply in Adams and Haynes (1986, 0.917) markedly differs from that of Myneni et al. (1994, 0.19). As such, the measured welfare changes and the recommended policy options may suffer from a lack of precision and pertinence.

Second, certain data deficiencies exist. In demand side, it is more appropriate to include personal disposable income, instead of GDP or GNP (Wear and Lee 1993, Lindsey et al. 2000). One major reason is that demand for lumber is mainly attributed to demand for new housing and repair & remodeling. The former is stimulated by housing starts, while the latter is driven by disposable income (USDA Forest Service 2001). These two variables are thus convenient measures of the likely impact of a stronger economy on lumber consumption (Yin and Baek 2004). Further, unlike disposable income, the effect of GDP or GNP on lumber price and/or consumption in previous studies is questionable. For example, such studies that used GDP or GNP, as an alternative to disposable income found that the estimated coefficients were consistently insignificant (Wear and Lee 1993, Lindsey et al. 2000). In contrast, the studies that employed disposable income found that the estimated coefficients were significant (Myneni et al. 1994, Sarker 1996). This implies that econometrically, disposable income is a more relevant variable for explaining the demand function (Wooldridge 2000). In addition, lumber demand is mainly attributed to housing demand, which is in turn strongly affected by financing trends, especially mortgage rates, instead of the 3-month T-bill interest rate (Jennings et al. 1994).

In supply side, previous studies paid little attention to log prices/markets. Logs are a key input in lumber production, accounting for about 70% of variable lumber production costs (Ragosta and Clark 2000). Excluding log prices in the supply function thus may cause the problem of biased estimates in assessing the impacts of the trade restrictions. This is called the omitted variable bias (Wooldridge 2000). In addition, log and lumber prices tend to track each other closely over time (Yin 2001). Moreover, since the 1990s, the federal harvest reductions in the Pacific Northwest (PNW) and the Asian financial crisis have contributed to a period of more volatile log and lumber prices. Without accounting for the log market dynamics, therefore, previous estimates of the trade restriction impacts might have not been robust (Yin and Baek 2004).

In addition, it is necessary to include lumber processing capacity and sawmilling wage rates in the supply function. First, given its high correlation coefficient with lumber production (0.7), lumber processing capacity can be represented as a proxy for capital costs in the supply function (Zhang 2001). Second, since lumber is a direct and primary output of sawmill industry, it is more reasonable to adopt sawmilling wage rates, instead of construction wage rates (Sarker 1996), as labor costs in the supply function.

In sum, there have been a number of studies on the impacts of the restrictions on imported Canadian lumber. However, due to their weaknesses in modeling approaches and variable uses, appropriately constructed and complete analyses are few. As such, with the intensified dispute and search for solutions, there remains a strong need for improved and up-to-date analysis of the lumber trade and quantification of the market and welfare effects of the U.S. trade restrictions.

For a more careful and comprehensive analysis, it is crucial to analyze the softwood lumber market structure appropriately and then to assess the effects of the trade restrictions and other market factors on lumber prices, domestic production and imported volume accurately. It is because all of the welfare measurements depend on these prices and volume effects from the trade restrictions. The estimated welfare consequences in turn indicate the effectiveness of a specific trade restriction (Yin and Baek 2004).

Therefore, this study assesses the impacts of the trade restrictions with improved variables and models. First, this study expands the list of variables, including variables from the log markets in the U.S. and Canada. Given that the central argument of the trade dispute has revolved around Canadian stumpage subsidies, it is essential to incorporate log prices when examining the price, volume, and welfare effects of the trade restrictions. In so doing, this study can properly decompose the effects of the trade restrictions and other market variables on the domestic lumber market.

Second, this study uses both the structural SSE model and the nonstructural VAR model in the assessment. This approach allows us not only to assess the market and welfare impacts of the trade restrictions in a corroborative way, but also to take advantage of both systems. Specifically, the VAR model enables us to explore the dynamic linkages of economic variables associated with the U.S.-Canada lumber trade without imposing *a priori* theoretical structure. Also, this model can capture short run and long run responses of an endogenous variable to other variables endogenous to the system, and reveal the potential effects of a trade restriction on every endogenous variable. It should be noted that so far, few studies have used a VAR model to assess the trade restriction effects and the dynamic interactions of the related market variables at the same time. On the other hand, a structural SSE model allows us to obtain the price elasticities, so that the welfare effects of the trade restrictions can be directly calculated. Moreover, with an expanded list of variables, these models will make it possible to contrast our estimates with those of previous studies. These complementary features should allow us to draw more balanced and robust conclusions.

Finally, this study discusses the potential effects of different restriction regimes. The main focus will be on what the most feasible and effective policy options are, how the gross as well as distributional welfare effects vary under different regulating regimes, and what strategies the U.S. government and industry should adopt in negotiating with their Canadian counterparts. In so doing, this study not only provide some estimates of what gains or losses to both countries are under different trade restrictions, but also examine the practicality of the restriction options.

1.2. Objectives

The objectives of this study are:

(1) To analyze the U.S. softwood lumber market structure;

(2) To assess the impacts of the MOU and the SLA;

(3) To examine the welfare and policy implications of different U.S. restriction regimes.

The specific questions that this study will answer are:

(1) How have the trade restrictions and market factors affected lumber prices, domestic production and imported volume?

(2) What are the welfare implications of these prices and volume effects?

(3) What are the most feasible and effective policy options for the U.S. to deal with the lumber trade?

1.3. Hypothesis to Be Tested

The hypothesis to be tested is: the U.S. trade restrictions have caused the U.S. lumber prices to increase, imported volume to decrease, and the U.S. lumber production to increase more than would have occurred without the trade restrictions. This study will test this hypothesis using both the structural SSE model and the nonstructural VAR model with an expanded list of variables.

1.4. Dissertation Outline

To accomplish the objectives and answer the specific questions, my dissertation consists of six chapters in addition to the introduction.

Chapter 2 provides an overview of the North American lumber market and the trade dispute between the U.S. and Canada over the last two decades. The overview includes historic trends of consumption, production, and trade in the U.S. and Canada, and their causal relationships with other market factors such as housing starts and exchange rates. A description of the core issues and the history of the trade dispute are then presented.

Chapter 3 presents data description and testing for unit roots. Particularly, this chapter provides extensive discussion on unit root tests. One of the major reasons is that it is prerequisite to test for the existence of unit roots in time-series data for estimating an appropriately specified VAR model adopted in chapter 5. Moreover, it is necessary to include tests for structural change since standard tests of unit roots, including the augmented Dickey-Fuller (ADF) test have lower power and are unable to detect a structural break in the series. After a description of the data, tests for structural change with known and unknown break points are performed, and then the stationarity of the time-series data are discussed.

Chapter 4 provides empirical analyses of the structural models — a simultaneous system of equations (SSE) and a reduced form equation (RFE) — for the U.S. softwood lumber market. The main tasks of the SSE model are to assess the price and volume impacts of a restrictive regime and to obtain the price elasticities used for welfare analysis in chapter 6. The RFE model is adopted to estimate the price and Canadian export impacts of the restrictions and to detect the potential variations between this and other studies. The theoretical framework and economic models as well as the estimation procedures for the structural model are discussed. A discussion of the empirical results and comparisons with other studies are also presented.

Chapter 5 provides empirical analyses of the nonstructural models — a vector auto-regression (VAR) and an error-correction (VEC) model. The main tasks of the VAR/VEC models are to not only discern the dynamic linkages among related market variables but also assess the impacts of the U.S. restrictions in a corroborative way. The theoretical framework for the Johansen co-integration test and the VAR/VEC models are discussed. The empirical results as well as the estimation procedures for the two models are discussed, followed with a summary and conclusions.

Chapter 6 includes the welfare impacts of the trade restrictions and simulation for different future policy options. The theoretical framework and the welfare consequences for the MOU and the SLA are presented. The simulated results for various policy options are then discussed, followed with a summary and conclusions.

Chapter 7 is a conclusive chapter that summarizes the major findings from the previous chapters. The implications of the results are also discussed and areas of future research are identified.

CHAPTER 2

U.S.-CANADA SOFTWOOD LUMBER TRADE

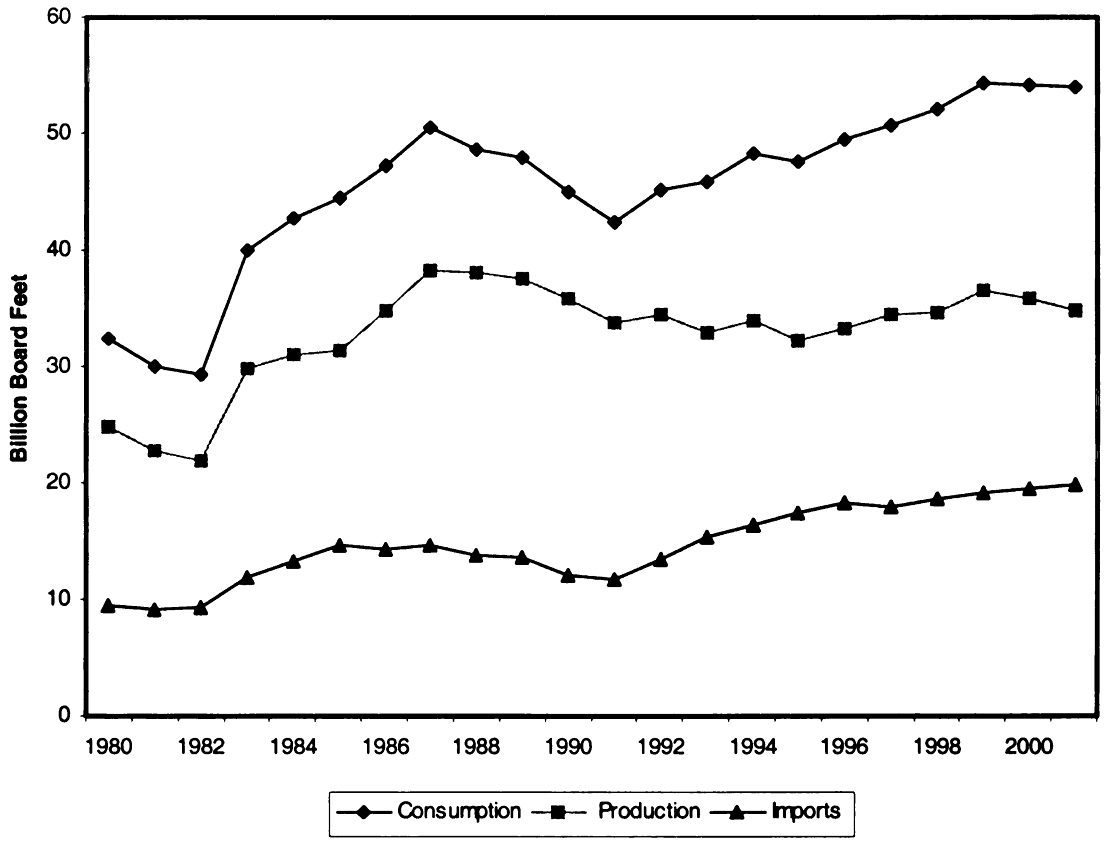
This chapter provides an overview of the North American softwood lumber market and the trade dispute between the U.S. and Canada over the last two decades. The overview includes historic trends of consumption, production, and trade in the U.S. and Canada, and their causal relationships with market factors such as housing starts and exchange rates. A description of the core issues and the history of the trade dispute are then presented, followed with a summary and conclusions.

2.1. North American Softwood Lumber Market

The trade of softwood lumber between the U.S. and Canada is one of the major trade flows of forest products in the world. These two countries are not only the world's largest producers of softwood lumber, but also the world's largest importer and exporter, respectively (Fukuda 2001). In 2000, the U.S. shared 45% of world imports for softwood lumber, while Canada provided 48% of world lumber exports (FAO 2001). More importantly, these two countries are highly interdependent on each other's lumber markets. That is, more than 90% of U.S. softwood lumber imports comes from Canada. Canada exports more than 60% of its softwood lumber production to the U.S. As a result, U.S. imports of softwood lumber have a strong influence on the North American lumber market.

2.1.1. U.S. Softwood Lumber Market

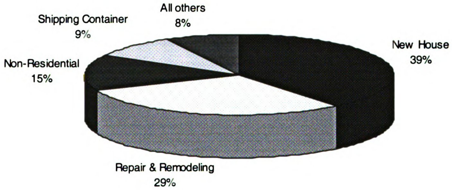
Figure 2.1 shows the levels of U.S. softwood lumber consumption, production, and imports for the period of 1980 to 2001. It can be seen that while U.S. consumption for softwood lumber has increased, the domestic production has not been consistent with the consumption. As a result, imports for softwood lumber have increased and thus provided a large percentage of the U.S. total consumption.



Data source: Statistics for the Wood Product Industry (various years)

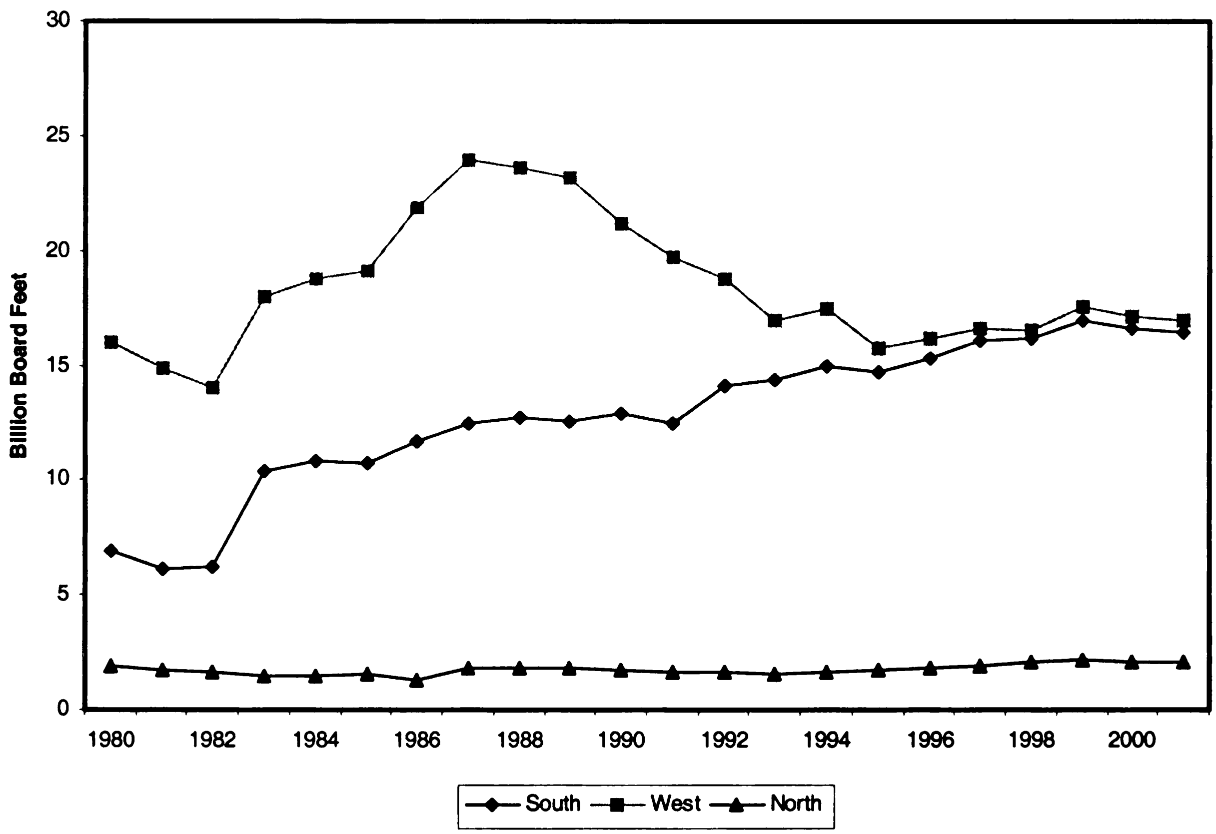
Figure 2.1. U.S. softwood lumber consumption, production, and imports (1980-2001)

During the last two decades, consumption of softwood lumber in the U.S. has consistently risen. Growing lumber consumption in the U.S. has mainly been attributed to an increase in housing demand, which consists of new houses and repair & remodeling (AF&PA 2001). As shown in Figure 2.2, in 2000 the share of these two sectors represented about 70% of the U.S. total consumption. In particular, with a strong housing market stimulated by unprecedented economic expansions, average annual lumber consumption in the U.S. rose to 52.2 BBF for the period of 1996 to 2000, which far exceeded the record high of 50.5 BBF in 1987.



Data source: AF&PA (2001)

Figure 2.2. U.S. softwood lumber consumption (2000)



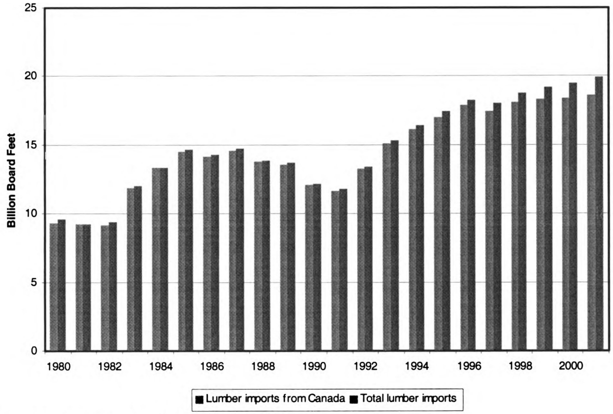
Data source: Statistics for the Wood Product Industry (various years)

Figure 2.3. U.S. softwood lumber production by regions[2] (1980-2001)

On the other hand, over the last two decades, softwood lumber production in the U.S. has not been consistent with growing lumber consumption. Specifically, in the 1980s U.S. lumber production had generally taken an upward trend, with all-time high levels of 38 BBF for the period of 1987 to 1989. In the early 1990s, however, the lumber production began to decline, due mainly to the timber harvest reductions from federal lands in the Pacific Northwest (PNW) region. As a result, the West, as the largest lumber

[2] The South includes Alabama, Arkansas, Florida, Georgia, Kentucky, Louisiana, Maryland, Mississippi, North Carolina, Oklahoma, South Carolina, Tennessee, Texas, and Virginia. The West includes Alaska, Arizona, California, Idaho, Montana, Oregon, and Washington. The North includes Illinois, Indiana, Iowa, Maine, Michigan, Missouri, New York, Ohio, Pennsylvania, and Wisconsin.

producing region in the U.S., has reduced its lumber production from 23.9 BBF in 1987 to 15.8 BBF in 1995, its share of U.S. production from more than 60% to about 50% (Figure 2.3). Consequently, the U.S. total production declined from 38.2 BBF in 1987 to 33.2 BBF in 1995. During the 1996-2001 period, due mainly to the increasing timber production in the South,[3] the lumber production has recovered up to about 34 BBF annually, close to that of the early 1990s.



Data source: Statistics for the Wood Product Industry (various years)

Figure 2.4. U.S. imports for softwood lumber (1980-2001)

[3] The South has increased its production from 12.5 BBF in 1987 to 14.7 BBF in 1995, increasing its share of U.S. production from 32.6% to 45.6% during the period. Currently, the South and the West region produce nearly equal amounts of softwood lumber (Howard 2000 and AF&PA 2001).

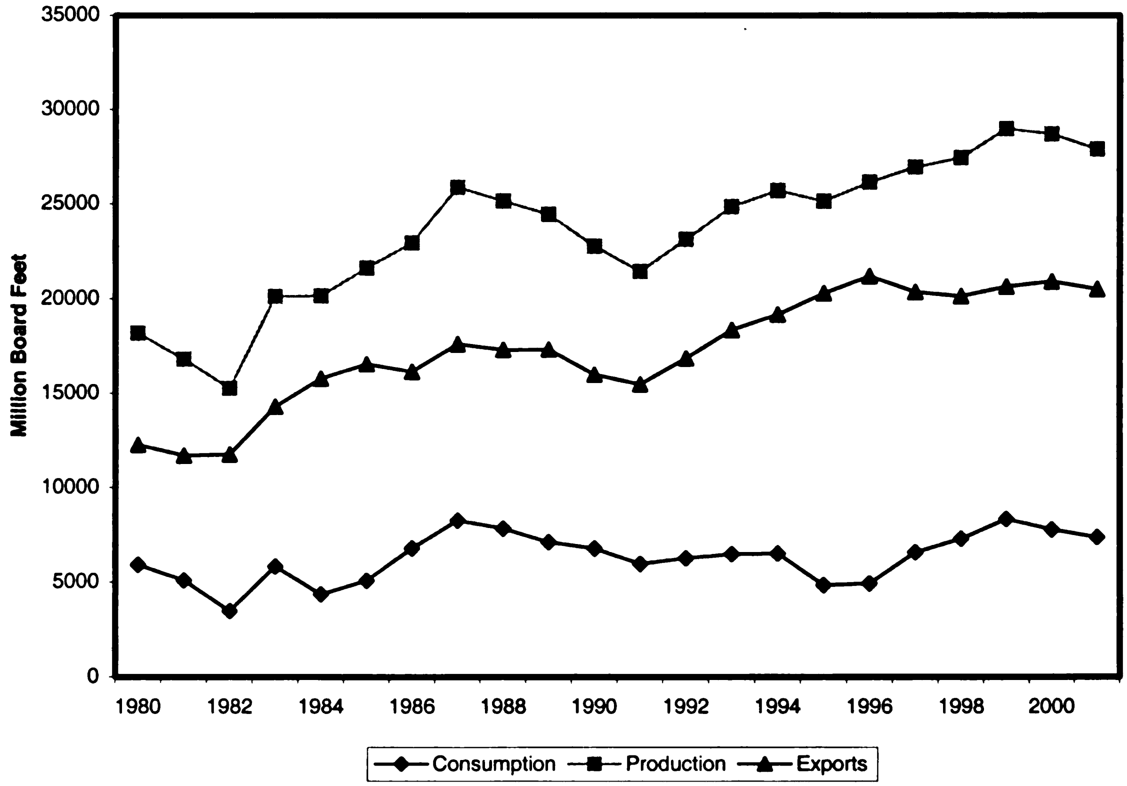
Coupled with the expanded domestic consumption and constrained domestic supply, the U.S. would not be able to satisfy its demand for softwood lumber without imports. As a result, during the last two decades, average annual imports of softwood lumber in the U.S. have risen from 12.4 BBF during the 1980-1990 period to 17.1 BBF during the 1991-2001 period (Figure 2.4). In particular, since 1992, import demand for lumber has increased substantially because of the timber harvest reductions in the PNW and the strong growth of the U.S. economy. In 2001, the total volume of softwood lumber imports reached a record high of 19.9 BBF.

Canada has remained the principal source of the U.S. softwood lumber imports, providing about 94% of total imports in 2000. Specifically, U.S. softwood lumber imports from Canada have grown fairly steadily in the 1990s. An average annual lumber import by the U.S. has risen to 16.5 BBF during the 1991-2000 period, up significantly from 12.3 BBF in the 1980s. Moreover, with the unprecedented expansion of the U.S. economy during the 1996-2001 period, U.S. lumber imports from Canada has reached a record high of more than 18 BBF annually, about 34% of total U.S. consumption for the period.

2.1.2. Canadian Softwood Lumber Market

Figure 2.5 shows the levels of Canadian softwood lumber consumption, production, and exports for the period of 1980 to 2001. In general, due to relatively small population in Canada, domestic consumption for softwood lumber has maintained steadily low over the last two decades. In contrast, Canada's production of softwood lumber has taken an upward trend. An average annual production of softwood lumber in

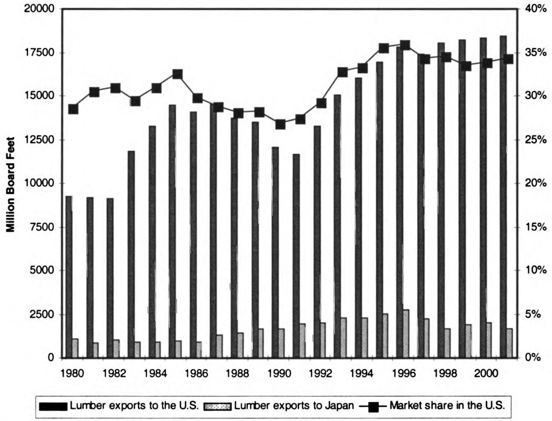
Canada has risen from 21.2 BBF during the 1980-1990 period to 26.0 BBF during the 1991-2001 period. Accordingly, an average annual export for softwood lumber in Canada has increased from 15.2 BBF during the 1980-1990 period to 19.5 BBF during the 1991-2001 period.



Data source: Statistics for the Wood Product Industry (various years)

Figure 2.5. Canada's lumber consumption, production, and exports (1980-2001)

A large percentage of Canadian production of softwood lumber has been exported to the U.S. over the last two decades. Particularly, about 63% of total production in Canada has exported to the U.S. during the 1991-2001 period, up from 58% in the 1980s. As a result, the Canadian share of U.S. market has reached from 29% in the 1980s to about 33% in the 1990s. Figure 2.6 shows the trends of Canadian exports of softwood lumber and market share in the U.S.



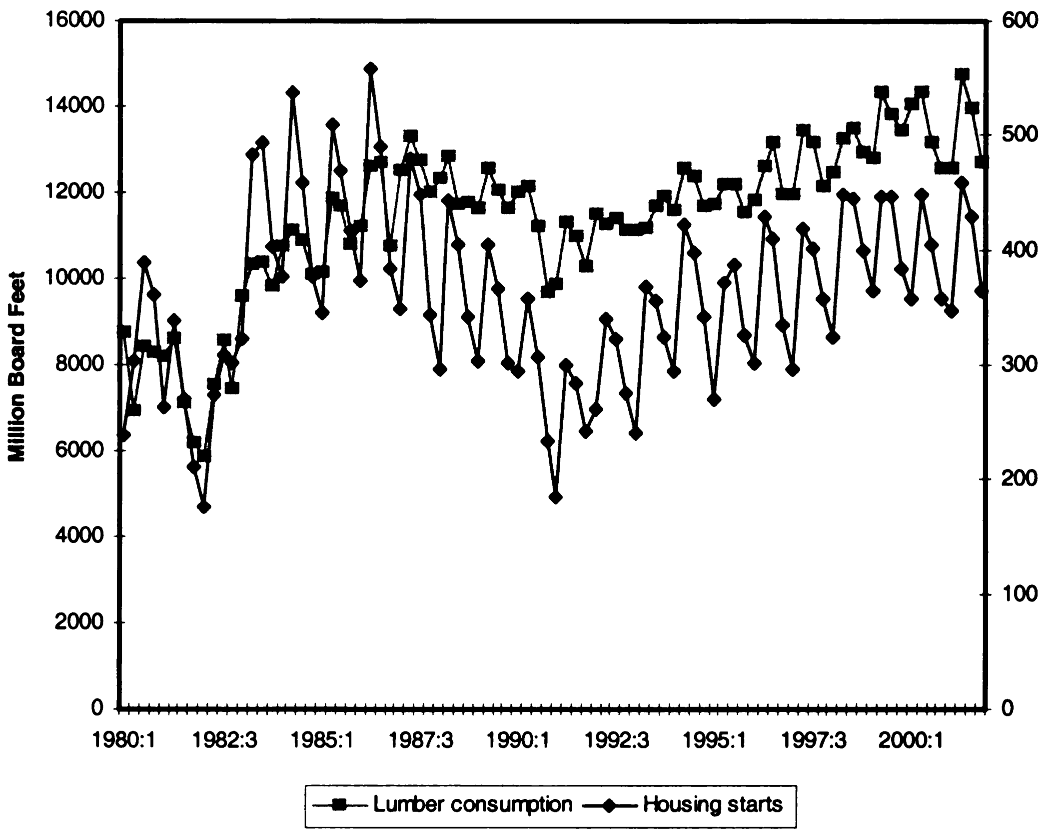
Data source: Statistics for the Wood Product Industry (various years)

Figure 2.6. Canada's lumber exports and market share in the U.S. (1980-2001)

2.1.3. North American Lumber Industry and Macroeconomic Variables

Figure 2.7 shows U.S. softwood lumber consumption and housing starts for the period of 1980 to 2001. The movements of the two series over the period are quite similar and thus show a strongly positive correlation (0.87). For instance, housing starts declined for the period of 1981 to 1982 and 1990 to 1991, respectively, due mainly to the economic recessions. In response, lumber consumption went down for those periods.

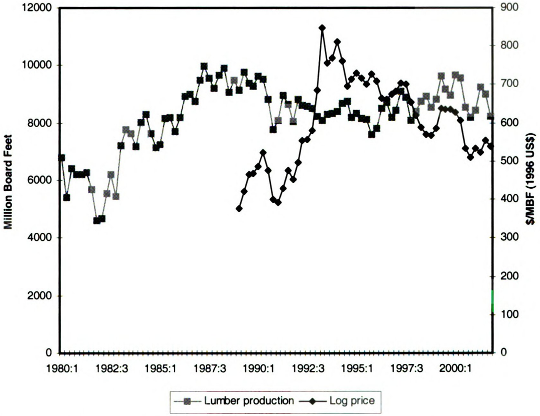
Since 1992, however, the U.S. economic boom has led to the consistent increase of housing starts, which in turn resulted in a steady increase of lumber consumption. In addition, lumber consumption is strongly correlated with personal disposable income (0.78) and mortgage rates (-0.84). However, lumber consumption is only weakly correlated with the 3-month T-bill interest rate (-0.26).



Data source: Statistics for the Wood Product Industry (various years)

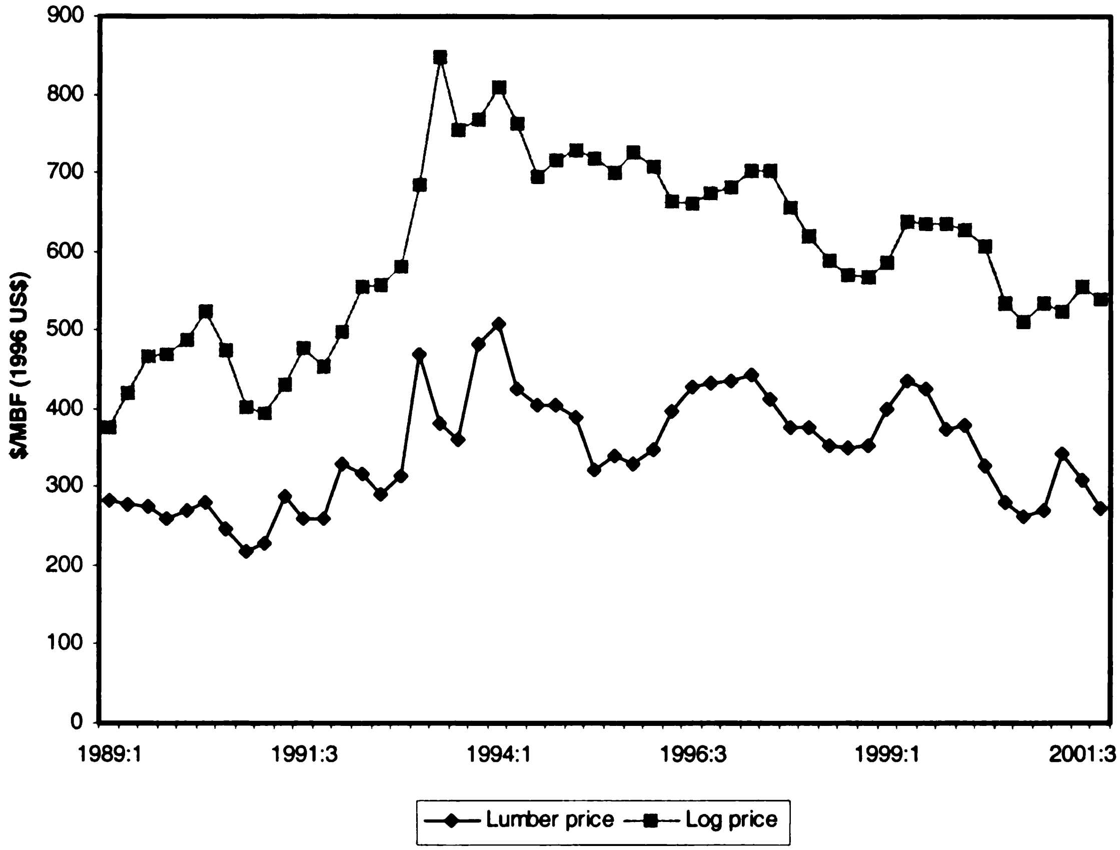
Figure 2.7 Softwood lumber consumption and housing starts (1980:1-2001:4)

Figure 2.8 shows U.S. softwood lumber production for the period of 1980 to 2001 and the price for Douglas-fir sawlogs (#2) in the PNW for the period of 1989 to 2001. The U.S. production has been strongly correlated with the log price (-0.78) over the last decade. As an example, in the early 1990s, due to the federal harvest reductions, production in the PNW, the largest timber production region, has been reduced significantly. As a result, the U.S. total production declined and the log price increased considerably. On the other hand, in the late 1990s, due mainly to the increased capacity in the South, lumber production has recovered nearly to the volume of the early 1990s. Consequently, the log price has drifted down over the period.



Data source: Statistics for the Wood Product Industry (various years)

Figure 2.8. U.S. softwood lumber production and log prices (1980:1-2001:4)

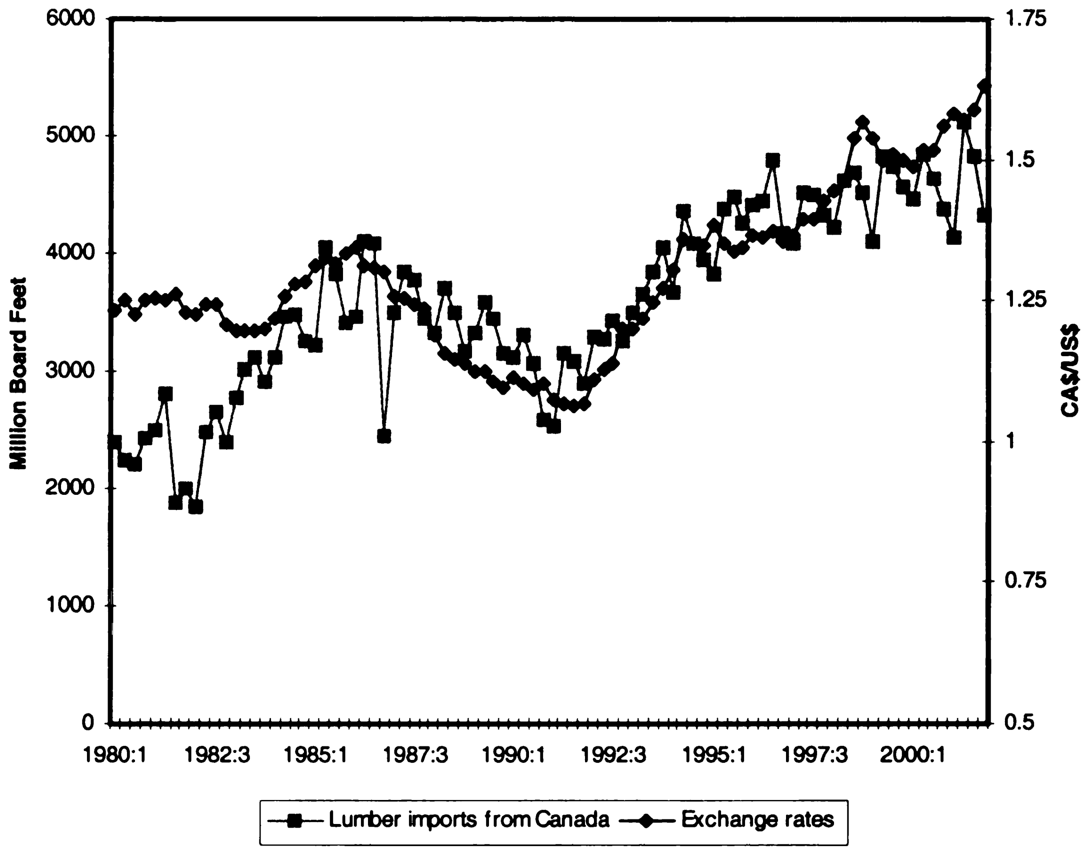


Data source: Random Lengths (2002) & Log Line (2002)

Figure 2.9. U.S. softwood lumber price and log prices in the PNW

Figure 2.9 shows U.S. softwood lumber prices and log prices in the PNW for the period of 1989 to 2001. The two price series tend to track each other closely over time and thus show a strongly positive correlation (0.83). For example, due to the federal harvest reductions in the PNW, the average log price went up from $380/MBF in 1989 to $580/MBF for the period of 1991 to 1994. During the same period, the average lumber price increased to $332/MBF, up from $240/MBF in 1989. In contrast, since late 1997, the Asian financial crisis has sharply reduced U.S. log exports to Asian countries, particularly Japan, which thus brought the average log price down to $490/MBF for the period of 1997 to 1998. Consequently, the average lumber price has dropped to

$295/MBF for the same period. Strong linkages thus exist between the lumber and timber markets in the U.S.



Data source: IMF & Statistics for the Wood Product Industry (various years)

Figure 2.10. U.S. lumber imports from Canada and exchange rates (1980:1-2001:4)

Figure 2.10 shows the U.S. softwood lumber imports from Canada and the real exchange rate between the U.S. and Canada for the period of 1980 to 2001. Softwood lumber imports from Canada have a strongly positive correlation (0.9) with the exchange rate. This indicates that exchange rate is one of the most significant factors in explaining the increased quantity of U.S. lumber imports from Canada over the last two decades. For instance, the U.S. dollar has strengthened against the Canadian dollar by nearly 30% for the period of 1992 to 2001. If all other factors were held constant, such an appreciation of the U.S. dollar against the Canadian dollar could represent a considerable improvement in the competitive advantage of Canadian lumber producers in the U.S. market. Indeed, Canadian lumber exported to the U.S. has increased to about 40% in close parallel with rising exchange rates over the period.

2.2. U.S.-Canada Softwood Lumber Trade Dispute

2.2.1. Core Issues of the Trade Dispute

The softwood lumber trade dispute between the U.S. and Canada extends back over the last 20 years. At the core of the dispute is the U.S. claim that Canada subsidizes its softwood lumber industry with low stumpage rates — fees paid by forest products companies to the provinces for the rights to cut trees. This allegation is based on the different systems of forestland ownership and different methods of stumpage pricing in the two countries.

In Canada, the provinces are mainly responsible for forest management. As shown in Table 2.1, most of Canada's forests (94%) are publicly owned (known as Crown lands), with 71% managed by provinces and 23% managed by the federal government. Specifically, British Columbia is the province holding the most valuable timberland, where about 96% of timberland is publicly owned. In addition, about 90% of forestlands is publicly managed in Quebec, Ontario, and Alberta. Notice that most of the 6% private forestlands is found in Manitoba and Nova Scotia (Statistics of Canada 2001).

Public forest policy in Canada has been designed to make timber available in ways that would support the development of increased processing capacity, and thus contribute to regional economic development (Irland 1986). To this end, provincial governments sell the stumpage on Crown lands to private forest companies under long-term leasing or licensing arrangements, known as the tenure system. In exchange, forest companies are required under these arrangements to create jobs and maintain community economic bases (Cashore 1998). In this circumstance, the provincial governments have determined stumpage fees administratively, rather than on the basis of market forces.

In contrast, only 42% of forestlands in the U.S. is publicly owned, with 23% managed by states and 19% managed by the federal government, whereas 58% of timberlands is privately owned. In particular, the South is dominated by private forestlands, where about 90% of timberland is privately owned. In addition, about 75% of timberland in the North is privately owned. Public ownership is predominant in the West, where about 70% of timberland is publicly owned.

Public forest policy in the U.S. has mainly evolved to reflect multiple interests including timber harvests, environmental preservation, and recreational use. Hence, regional economic development is generally in a secondary position behind other values (Irland 1986). Moreover, much of the timber in the U.S. sold by public and private landowners are based on competitive bidding systems. Therefore, the U.S. lumber industry argues that Canadian timber from Crown lands is sold at a lower stumpage rate administratively set by provincial governments. As a result, these lower stumpage rates include a subsidy to the Canadian lumber industry.

Table 2.1. Ownership of forestlands in the U.S. and Canada (%)

Ownership	U.S. (1997)				Canada (2001)					
	North	South	West	Total	BC	Quebec	Ontario	Alberta	Maritimes	Total
Private	75	88	31	58	4	11	11	4	58	6
State	18	6	35	23	95	89	88	87	40	71
Federal	7	6	34	19	1	0	1	9	2	23

Data Source: USDA Forest Service (1997) and Statistics of Canada (2001)
Note:
1. See footnote 1 for the definitions of North, South and West in the U.S.
2. State in the U.S. corresponds to province in Canada.
3. BC represents British Columbia.

2.2.2. History of the Trade Dispute

2.2.2.1. The First Countervailing Duty Investigation

In 1981, the Northwest Independent Forest Manufacturers (NIFM) first alleged that Canadian subsidized softwood lumber resulted in increased unemployment in the Pacific Northwest (PNW) forest industry. The Senate Finance Committee asked the International Trade Commission (ITC) to investigate this complaint. In April 1982, the ITC found that Canadian lumber producers gained market share with their low timber costs. However, the ITC report made no recommendation, and did not constitute a finding of injury (Cashore 1998).

In July 1982, NIFM enlisted the support of companies outside of PNW and formed the nationwide Coalition for Fair Canadian Lumber Imports (CFCLI, now the Coalition for Fair Lumber Imports, CFLI). In October 1982, the CFCLI formally launched its first countervailing duty (CVD) petition against Canadian softwood lumber imports, known as Lumber I.[4] The CFCLI alleged that Canadian stumpage rates provided a subsidy and materially injured U.S. lumber producers. Investigation by the Department of Commerce (DOC) and the ITC focused on the stumpage programs in British Columbia (BC), Alberta, Ontario, and Quebec. In November 1982, the ITC ruled that there is a "reasonable indication" that the U.S. was being injured by Canadian lumber imports. In

[4] A countervailing duty (CVD) is a special duty imposed to protect domestic industry from injury caused by imports that have benefited from subsidies provided by a foreign government. A CVD can only be applied if it has been established in an investigation that imported goods have been subsidized and that such subsidized imports are causing material injury, or threatening to cause injury to the domestic industry. According to the U.S. trade law, the ITA in the DOC must determine whether a foreign government is directly or indirectly providing a countervailing subsidy for the manufacture, production or exportation of merchandise imported or sold into the United States. On the other hand, the International Trade Commission must determine whether the U.S. industry producing the like goods has been materially injured, or threatened with material injury, by reason of subsidized imports. If these two conditions are met, a CVD equal to the amount of the subsidy is imposed upon the imports of the subsidized merchandise (Department of Foreign Affairs and International Trade 2001).

May 1983, however, the International Trade Administration (ITA) in the DOC concluded that stumpage practices did not automatically restrict timber "specially" to the lumber industry and stumpage prices did not discriminate in an unfair manner by providing timber to exporters at preferential rates (Hoberg and Howe 1999).

2.2.2.2. The Second Countervailing Duty Investigation

In May 1986, the U.S. producers again petitioned for another CVD investigation with the DOC and the ITC. This is referred to as Lumber II. The U.S. lumber industry requested a tariff of 27% on imports of softwood lumber from Canada. The ITC again found preliminary evidence of injury to the U.S. lumber industry. In October 1986, the ITA reversed its 1983 determination and ruled that Canadian producers received a subsidy of 15% *ad valorem* (i.e., 15% of lumber prices). The ITA explained that court rulings, legislative changes and the correction of errors it made in 1983 were the causes of this reversal (Cashore 1998).[5] On December 30, 1986, the day before the final ITA determination, the U.S. and Canada negotiated an agreement (the Memorandum of Understanding, MOU), with a 15% tax imposed on Canadian lumber exported to the U.S. That export tax went into effect on January 8, 1987.

Subsequent amendments to the MOU allowed provincial governments to reduce or eliminate the 15% export charge by implementing so called "replacement measures", defined as increased stumpage or other provincial charges on softwood lumber production. British Columbia changed their stumpage pricing system in 1987 and the

[5] In 1984, the U.S. Congress amended the Tariff Act of 1930 to include "upstream" or "input product" subsidies as countervailable where "the input product bestows a competitive benefit on those goods by affecting significantly the cost of production" (Tougas 1988-89, p.144).

U.S. government agreed to the elimination of the 15% export tax. In addition, Quebec also changed their stumpage prices, allowing the province to reduce the export tax to 3.1%.

2.2.2.3. The Third Countervailing Duty Investigation

In October 1991, the Canadian government unilaterally declared the termination of the MOU, stating that it had sufficiently increased stumpage fees to eliminate the export tax. In response, the US Trade Representative (USTR) immediately imposed provisional duties on Canadian lumber. Specifically, a 6.2% duty was set on Quebec, while a 15% duty is placed on softwood lumber from Ontario, Alberta, Manitoba, and Saskatchewan. Softwood lumber from British Columbia was exempted since higher stumpage fees imposed as a result of the MOU remained in effect. Meanwhile, the DOC self-initiated a new CVD investigation without a petition from lumber industry and imposed a temporary bonding requirement on lumber imports. This was termed Lumber III — the first CVD case after the signing of the U.S.-Canada Free Trade Agreement (FTA).

In March 1992, the ITA issued a preliminary finding of 14.48% *ad valerom* subsidies. With a final determination, in May 1992, the ITA established a 6.51% *ad valerom* subsidy — 2.91% for stumpage and 3.6% for log export restraints in British Columbia. As a result, a CVD of 6.51% was imposed on lumber imports from the major four provinces (British Columbia, Ontario, Quebec and Alberta), Saskatchewan, Manitoba, the Yukon Territory and the Northwest Territories. Notice that in addition to investigating the traditionally alleged subsidies of Canadian stumpage fee, the U.S.

government targeted another Canadian forestry practice, which is log export restrictions in British Columbia.[6] The claim was that these restraints served to lower domestic log prices in British Columbia, which provided a subsidy to its lumber producers and other firms using unprocessed logs as an input to their manufacturing processes (Ragosta and Clark 2000).

In August 1992, Canada immediately responded to these rulings by appealing to the General Agreement on Tariff and Trade (GATT) and to binational panels established under Chapter 19 of the U.S.-Canada FTA. The GATT panel ruled that the U.S. acted properly when it self-initiated the CVD investigation. However, the bilateral panel found that the US acted improperly when it imposed interim import bonds on Canadian shipment to the US. In May 1993, the panel remanded the ITA finding for further analysis. In September 1993, the ITA revised its findings from 6.51% to 11.54% *ad valerom* subsidies. In December 1993, the panel again remanded the ITA finding and ordered the ITA to find no subsidies. In January 1994, the ITA announced that, in accordance with the panel's instructions, it determined that Canada's stumpage and log export policies were not countervailable.

In April 1994, the U.S. Trade Representative (USTR) requested the establishment of an Extraordinary Challenge Committee (ECC) under the Chapter 19 of the FTA to review the binational panel decisions. However, the ECC was dismissed for failing to meet FTA standards. In August 1994, the DOC published a notice in the Federal Register confirming termination of the CVD order and declared liquidation of all entries. All countervailing duties tentatively paid by Canadian exporters (approximately $800

[6] The U.S. government made a conclusion that log export restrictions in Quebec, Alberta and Ontario had no practical effect (Cashore 1998).

million) were refunded. In December 1994, the U.S. Congress passed legislation that specifically changed/clarified those parts of U.S. trade law that the binational panels used in their rulings, trying to ensure that future binational panels would no longer misinterpret U.S. trade law.[7] This legislation threatened the Canadian government and industry with the possibility of another CVD action. Accordingly, Canada sought a compromise with the U.S., holding a number of formal and informal meetings. Finally, in May 1996, the U.S. and Canada signed the "Softwood Lumber Agreement between the Government of Canada and the Government of the United States of America (SLA)."

2.2.2.4. The U.S.-Canada Softwood Lumber Agreement

The Softwood Lumber Agreement (SLA) was signed on May 29, 1996 and went into effect retroactively from April 1, 1996, to March 31, 2001. The SLA is a voluntary export restraint (a tariff-regulated quota system) agreed upon by the Canadian government, so that only Canada imposed a governmental regulation on softwood lumber exports. In addition, the SLA had no limitation imposed on the volume of softwood lumber exported from Canada. Canadian producers could export as much softwood lumber as they liked into the U.S., as long as they were willing to pay the export fees for the quantity exceeding the base volume (Fukuda 2001).

The SLA imposed export regulation on softwood lumber exported from the four major softwood lumber producing provinces in Canada: British Columbia, Alberta, Ontario, and Quebec. Softwood lumber producers operating in the four provinces were

[7] In December 8, 1994, the Uruguay Round Agreements Act explicitly approved the President's "statement of administrative action (SAA)", which stated because of Canadian practices, lumber imports from Canada could be subject to a CVD (Gorte and Grimmett 2002).

permitted to ship 14.7 BBF of lumber a year to the U.S. duty-free, referred to as the "established base." The next 650 million board feet (MBF) — export volumes between 14.7 BBF and 15.35 BBF — were subject to a fee of $50/MBF, referred to as the "lower fee base." All greater quantities, in excess of 15.35 BBF, were subject to a charge of $100/MBF, referred to as the "upper fee base." The fee levels could be adjusted for inflation annually based on an average of Consumer Price Indices (CPI) in the two countries. The trigger price was defined as the average for a complete calendar quarter of the price reported by Random Lengths for Spruce-Pine-Fir, Eastern, Kiln Dried, 2×4, Standard & Better, delivered to the Great Lakes. If the trigger price met or exceeded $405/MBF in any quarter during the first two years or $ 410/MBF thereafter, 92 million board feet of additional tax-free exports would be allocated in the subsequent quarters within one year of issuance. According to the SLA, Canada allocated the export quota to the four provinces based on historical shipments to the U.S.: British Columbia (8.673 BBF, 59%), Alberta (1.132 BBF, 7.7%), Ontario (1.514 BBF, 10.3%), and Quebec (3.381 BBF, 23%). The volume of softwood lumber exported from the four provinces into the U.S. over the past five year (1996-2001) is summarized in Table 2.2.

On the other hand, as a part of the agreement, the U.S. government agreed not to pursue trade remedy actions under CVD or other trade laws for five years. However, several additional disputes (such as reclassification of pre-drilled studs, rougher-header lumber, and notched studs[8], and stumpage fee reduction in British Columbia) arose

[8] In 1999, the U.S. Customs proposed to reclassify pre-drilled studs, notched lumber and rougher-headed lumber from a classification in the Harmonized Tariff Schedule of the United States (HTSUS) not covered by the SLA into a category that was covered. Canada responded by petitioning the reclassification of drilled studs to the World Customs Organization (WCO). The WCO ruled in Canada's favor, but the ruling was non-binding. Finally, just before the expiration of the SLA, an arbitration panel ruled that the U.S. violated the SLA by reclassifying drilled studs and notched lumber (BC STATS 2001 May).

during the SLA period.[9] The SLA expired in March 31, 2001 without either an extension or a new agreement.

Table 2.2. Canadian Lumber Exports to the U.S. under the SLA

		British Columbia	Alberta	Ontario	Quebec	Total
April 1996-March 1997	Export (MBF)	8,858,275	1,209,841	1,821,716	4,019,866	15,909,700
	Share (%)	55.6	7.6	11.5	25.3	100.0
April 1997-March 1998	Export (MBF)	8,738,284	1,218,272	1,665,270	3,935,347	15,557,174
	Share (%)	56.2	7.8	10.7	25.3	100.0
April 1998-March 1999	Export (MBF)	8,579,010	1,203,022	1,719,746	3,921,669	15,423,448
	Share (%)	55.6	7.8	11.2	25.4	100.0
April 1999-March 2000	Export (MBF)	8,681,546	1,324,227	1,728,025	3,945,664	15,679,465
	Share (%)	55.4	8.4	11.0	25.2	100.0
April 2000-March 2001	Export (MBF)	8,480,384	1,208,800	1,669,817	3,835,869	15,194,871
	Share (%)	56.0	7.7	11.0	25.3	100.0

Data source: Department of Foreign Affairs and International Trade (2001)
Note:

1. A calendar quarter for the SLA starts from April between 1996 and 2001.
2. MBF represents thousand board feet.

2.2.2.5. The Fourth Countervailing Duty Investigation

When the SLA expired in March 31, 2001, the U.S. softwood lumber industry immediately petitioned for a new CVD investigation. In addition to the allegation of subsidies to Canadian lumber producers, the U.S. industry, for the first time, also accused the Canadian industry of dumping — selling below costs — softwood lumber into the U.S. market. The U.S. producers required the imposition of 39.9% CVD and 28%-38%

[9] In May 1998, the BC government announced that it would lower stumpage fees in the province, effective June 1. Fees to Coast mills thus dropped an average of CA\$8.10/m^3 and fees to Interior mills dropped an average of CA\$3.50/m^3. The main reason of the reduction was that the Coastal forest industry suffered from a serious economic downturn, due to the Asian financial crisis. Specifically, Japanese markets plummeted while the SLA made it impossible for companies, especially along the Coast, to redirect shipments to the U.S. market without exceeding their individual quotas. The CFLI countered with a challenge under the SLA dispute settlement procedures. In July 1999, a settlement between the two countries was announced, which called for BC's upper- and lower-fee base shipments during the fourth and fifth years of the SLA to be set at their average for the first and second years. This resulted in the province's share of the 650 million board feet (lower-fee base) being lowered by 90 million board feet per year. The agreement also set an upper-fee base limit of 110 million feet. Accordingly, shipments exceeding that amount would carry a "super fee" of \$146.25/MBF in the fourth year and \$148.47/MBF in the fifth year (Department of Foreign Affairs and International Trade 2001).

anti-dumping duty (ADD) on imported Canadian lumber.[10] This episode is referred to as Lumber IV.

In April 2001, the DOC announced that it was initiating CVD and ADD investigation. The DOC decided to investigate first whether provincial and federal forest management regimes (including stumpage and log export bans) confer a countervailing duty. In August 2001, the DOC issued its preliminary determination with finding of 19.31% *ad valorem* subsidies. The DOC also made the affirmative critical circumstances determination, concluding that there had been a surge of softwood lumber exports from Canada since April 1, 2001. As a result of these two decisions, the DOC could apply retroactive measures, in the form of bonds or cash deposits in the amount of 19.31% to shipments made on or after May 17, 2001. Later, in October 2001, the DOC issued its preliminary dumping determination and found that producers/exporters of softwood lumber from Canada sold their product below market value. In contrast to the CVD case, the DOC preliminarily found that critical circumstances did not exist on the anti-dumping case. In March 2002, the DOC announced its final determinations of CVD and anti-dumping cases. In May 2002, with the final material injury determination ruled by the ITC, the DOC applied CVD (18.79%) and ADD (average 8.43%) on Canadian softwood

[10] Dumping is the sale of goods in foreign markets at prices that are below those charged for comparable sales in the home market or that are below the cost of producing the goods. Unlike a CVD that examines the subsidy practices of governments, anti-dumping investigation concern the pricing practices of individual firms. Under an anti-dumping investigation, only select companies would be investigated to determine an "all others" rate. According to the U.S. trade law, the ITA in the DOC must determine dumping margins by comparing the price at which the subject goods are sold in the U.S. with the "normal value" of the goods. Normal value is the price at which comparable sales of the subject goods are made in the home market. If the volume of home market sales is insufficient or if such sales are not made in the ordinary course of trade, normal value can be determined on the basis of sales to a third country, or on the cost of producing the goods, including a reasonable amount for profit. On the other hand, the ITC must determine whether the U.S. domestic industry producing like products is materially injured, threatened with material injury or, the establishment of an industry in the U.S. is materially retarded by reason of the subject imports (Department of Foreign Affairs and International Trade 2001).

lumber. In responses, Canada has filed the CVD and ADD cases to the North American Free Trade Agreement (NAFTA) and World Trade Organization (WTO).

In January 2003, the DOC released a policy bulletin, "Proposed Analytical Framework, Softwood Lumber from Canada", known as Aldonas Proposal, which provided the procedure to bring to an end the CVD. That is, the DOC reviews the elimination of the 18.79% CVD on a province-specific basis, if a province in Canada changes its stumpage system to a market-based system sufficiently. Since then, with exchanging proposals, a number of negotiation sessions between the U.S. and Canada have been held to resolve the softwood lumber dispute. However, these have not resulted in a resolution of the dispute until the end of 2003.

In December 2003, the U.S. and Canada finally reached a tentative agreement to end the recent dispute. The proposed agreement eliminates the 27% U.S. tariff on Canadian lumber, but reduces the amount of softwood lumber that Canadian companies can export to the U.S. without penalty. That is, Canadian producers should drop their exports to 31.5% of U.S. total lumber consumption, and pay a tariff of $200 for each additional thousand board feet, which is referred to as "quota allocation principles." This deal also includes the 52% return of the $1.6 billion U.S. duties that Canadian companies have paid, with U.S. companies keeping the rest.

In the meantime, the WTO and the NAFTA released their decisions on Canada's challenge of the CVD and the ADD cases. In July 2002, the WTO panel released the preliminary ruling that the U.S. should abolish or change the Byrd Amendment, which stipulates the distribution of duties collected from the anti-dumping and countervailing duties to U.S. petitioners. In addition, the panel ruled that Canadian stumpage prices can represent a subsidy, although the U.S. government cannot use cross-border comparisons to determine the level of subsidy. In July 2003, after reviewing the ADD determination, the NAFTA remanded it back to the DOC to recalculate the 8.4% duty, which lowered to 8.1% on October. In August 2003, the NAFTA panel ruled on Canada's challenge of the CVD case, concluding that the DOC may not calculate the countervailing duties by comparing Canadian stumpage rates to U.S. stumpage rates. In January 2004, the WTO Appellate Body reversed the decision of July 2002, ruling that the DOC may use a benchmark other than private prices within Canada in "very limited circumstance" for estimating the magnitude of timber subsidies conferred to Canadian lumber industry. In addition, in April 2004, the WTO panel released its final decision on the ADD case, which found that the U.S. could impose anti-dumping duties on lumber imported from Canada. However, the WTO rejected the use of zeroing as a method for calculating anti-dumping rates and remanded the case back to the DOC to recalculate the duties.[11]

[11] Zeroing means the practice of assigning a margin of zero to goods for which the export price exceeds the home market price. Zeroing prevents the negative margin for one category of goods from offsetting a margin of dumping for another category of goods, thus inflating the overall dumping margin (Department of Foreign Affairs and International Trade 2001).

2.3. Summary and Conclusions

Over the last two decades U.S. consumption for softwood lumber increased, while the domestic production has not been consistent with the consumption, resulting in a consistent increase of lumber imports from Canada. As a result, with an increase of Canada's share of the U.S. lumber market, the two countries have involved in the softwood lumber trade dispute. The core issue of the dispute is the U.S. claim that Canada subsidizes its lumber industry with low stumpage rates. This is mainly attributed to the different systems of forestland ownership and thus different methods of stumpage pricing in the two countries. This allegation has led to four rounds of CVD cases. In 1982, U.S. producers filed their first CVD case against imported Canadian lumber, but no subsidies were found (Lumber I). In 1986, however, preliminary subsidy findings led to the MOU (Lumber II). With the unilateral termination of the MOU by the Canadian government in 1991, the U.S. imposed a 6.51% CVD on Canadian lumber (Lumber III). In 1996, the U.S. and Canada agreed to enter into the SLA. With the expiration of the SLA in 2001, U.S. producers filed CVD and anti-dumping cases against Canadian softwood lumber (Lumber IV). In December 2003, the U.S. and Canada reached a tentative agreement to end the recent dispute. This agreement would allow shipping to the U.S. fee free until it reached 31.5% of U.S. total consumption, and pay a fee of $200 for each thousand board feet.

CHAPTER 3

DATA DESCRIPTION AND TESTING FOR UNIT ROOTS

This chapter describes data used in this study and conducts diagnostic tests, especially unit roots. One major reason is that testing for the presence of unit roots in time-series data is a prerequisite for estimating an appropriate vector auto-regression (VAR) model adopted in chapter 5. In other words, the procedures to construct an appropriate VAR system are closely related to between the time-series and stationarity. For example, if all the variables are stationary or non-stationary in the general VAR system with n variables, it is appropriate to use unrestricted VAR system in level or first difference forms (Harris 1995). Moreover, since standard tests of unit roots, including the augmented Dickey-Fuller (ADF) test, have lower power and are unable to detect structural breaks in the series, it is thus necessary to include tests for structural changes of the series. After a description of the data, therefore, this chapter performs tests for structural changes with known and unknown break points to examine whether there is any evidence of structural breaks in the series under consideration (Perron 1989 and Banerjee et al. 1992). Unit root tests of the time-series data are then discussed in the last section.

3.1. Data Description

Data used in this study are quarterly observations for the period of 1980-2001. The data sources are as follows. U.S. lumber consumption, production and imports from Canada (in million board feet, MMBF) were collected from the Statistics for the Wood

Product Industry published by the American Forest and Paper Association (AF&PA). Lumber prices ($/thousand board feet, MBF) — framing lumber composite price (a weighted average of 15 key framing lumber prices) — were from Random Lengths. Log prices ($/MBF) in the PNW — Douglas fir (#2 sawmill logs) in Washington and Oregon — were taken from Log Lines. Housing starts (in thousand units) came from the U.S. Census Bureau. Personal disposable income (U.S.$) was from the Bureau of Economic Analysis, Department of Commerce. The 30-year fixed mortgage rate was taken from the U.S. Federal Reserve Bulletin. U.S. lumber processing capacity (in MMBF) came from the USDA Forest Service. Wage rates ($ per hour) in U.S. sawmill industry (Standard Industrial Classification, SIC, 242) were taken from the U.S. Bureau of Labor Statistics.

Canadian log prices (CA$/$m^3$, an average rate for BC coast) were from the Revenue Branch of the BC Ministry of Finance. Canadian sawmilling wage rates (CA$ per hour) and lumber processing capacity (in MMBF) were taken from the Canadian Forest Service. The exchange rate (expressed as Canadian $ per U.S. $) was obtained from International Financial Statistics published by the International Monetary Fund (IMF). Finally, CPI and PPI for the U.S. were from the Bureau of Labor Statistics and the Survey of Current Business, respectively. CCI and PPI for Canada were obtained from Statistics Canada and the Bank of Canada.

The dummy for the MOU was set to one for the period of 1986:4-1991:3, and zero otherwise. Likewise, the dummy for the SLA was set to one for the period of 1996:2-2001:1, and zero otherwise. Producer Price Indices (PPI) were used to derive real wage, and log and lumber prices for the two countries, and the U.S. Consumer Price

Index (CPI) was used to deflate personal disposable income. Tables 3.1 and 3.2 summarize the data descriptions and preliminary statistics.

U.S. log prices in the PNW could only be traced back to the first quarter of 1989 (Arbor-Pacific Forestry Service 2001). This study thus compiled two datasets for the analysis. Dataset I contained every variable but the PNW log prices for 1980:1-2001:4. To capture shocks from log markets in the U.S., therefore, two dummy variables were introduced in dataset I. The first dummy (DUM_1) represented the federal harvest reductions in the PNW for the period of 1991:1-1993:4. The second dummy (DUM_2) covered the Asian financial crisis for the period of 1997:3-1998:4. Dataset II covered all the variables for 1989:1-2001:4. Also, because all series are available in original form, seasonally unadjusted, this study will consider seasonality explicitly in the model.

Table 3.1. Data Description

Data	Abbreviation	Measurement	Source
Lumber Demand	DEM	MMBF[1]	AF&PA[3]
Lumber Production	SUP	MMBF[1]	AF&PA[3]
Canadian Lumber Export	EXP	MMBF[1]	AF&PA[3]
Lumber Prices	P	U.S.$/MBF[2]	Random Lengths
Housing Starts	HS	1,000 units	Census Bureau
Disposable Income	DI	U.S.$ per capita	Bureau of Economic Analysis
Mortgage Rates	MR	%	U.S. Federal Reserve Bulletin
Log Prices	LP	U.S.$/MBF[2]	Log Lines
Processing Capacity	PC	MMBF[1]	U.S. Forest Service
Wage Rates	W	U.S.$ per hour	U.S. Bureau of Labor Statistics
Canadian Log Prices	CLP	CA$/cubic meter	British Columbia Forest Service
Canadian Processing Capacity	CPC	MMBF[1]	Canadian Forest Service
Canadian Wage Rates	CW	CA$ per hour	Canadian Forest Service
Exchange Rates	EX	CA$/U.S.$	International Monetary Fund
Time Trend	T	Integer	1980:1=1, 1980:2=2, etc
Memorandum of Understanding	MOU	1 or 0	Dummy (=1 for 1986:4-1991:3)
Softwood Lumber Agreement	SLA	1 or 0	Dummy (=1 for 1996:2-2001:1)
Federal Harvest Reduction	DUM_1	1 or 0	Dummy (=1 for 1991:1-1993:4)
Asian Financial Crisis	DUM_2	1 or 0	Dummy (=1 for 1997:3-1998:4)
U.S. Consumer Price Index	CPI	1996=100	U.S. Bureau of Labor Statistics
U.S. Producer Price Index	PPI	1996=100	Survey of Current Business
Canadian Consumer Price Index	CPI	1996=100	Statistics Canada
Canadian Producer Price Index	PPI	1996=100	Bank of Canada

Note:

1. Million board feet.
2. Thousand board feet.
3. American Forest & Paper Association.

Table 3.2. Summary Statistics

Variable	Observations	Mean	Standard Deviation	Minimum	Maximum
DEM	88	11,391.3	1,871.7	5,846.0	14,723.0
SUP	88	8,224.1	1,185.7	4,593.0	9,984.0
EXP	88	3,598.9	791.4	1,840.0	5,124.8
P	88	282.7	85.1	163.0	475.0
HS	88	361.5	78.2	176.7	558.4
DI	88	5,124.8	1,889.7	2,248.1	8,706.2
MR	88	10.2	2.9	6.8	17.7
LP	52	579.4	119.8	328.0	793.0
PC	88	10,548.5	749.1	9,110.6	11,843.5
W	88	9.5	1.5	6.4	12.3
CLP[1]	88	77.9	28.9	43.2	138.9
CPC	88	6,561.5	592.8	5,122.5	7,159.5
CW	88	15.2	2.8	9.1	19.1
EX	88	1.3	0.1	1.1	1.6
T	88	-	-	1	88
MOU	-	-	-	0	1
SLA	-	-	-	0	1
DUM_1	-	-	-	0	1
DUM_2	-	-	-	0	1

Note:

1. Factor for converting cubic meter (m^3) to thousand board feet (MBF) is 2.36. Thus, CA$ 77.9/m^3 is equivalent of CA$ 183.9/MBF.

3.2. Testing for Unit Roots

When dealing with time-series data, it is necessary to test for the presence of unit roots not only to avoid the problem of spurious regression but also to set up an appropriate vector auto-regression (VAR) system. First, if a series has a unit-root, then it is non-stationary. Unless this non-stationary series combines with other non-stationary series to form a stationary co-integration relationship, the estimation can falsely represent the existence of a meaningful economic relationship, which is known as spurious regression problem (Wooldridge 2000). Note that if two time-series data are co-integrated, they will move closely together over time and the difference between them will be stable or stationary. Second, the procedures to construct an appropriate VAR system are linked closely to between the time-series and stationarity. For instance, if all the variables are stationary in the general VAR system with n variables, it is appropriate to use unrestricted VAR system in level forms. Or, if all the variables are non-stationary, say integrated of order one, or I (1),[12] but show no co-integration relationships, then it is necessary to estimate unrestricted VAR system in first difference forms. However, if all variables are non-stationary and co-integrated, it is appropriate to treat the system as a vector error-correction (VEC) model (Harris 1995).

Even though there are several tests for the presence of unit roots, the (augmented) Dickey-Fuller (ADF) test (Dickey and Fuller 1979) has been more popular due to its simplicity (Harris 1995). However, since the early 1990s, many studies have pointed out that the ADF test has critical shortcomings (Perron 1989, Banerjee et al. 1992, Blough

[12] Integrated of order zero, or I (0) is defined as a stationary, weakly dependent time-series process. On the other hand, integrated of order one, or I (1) is defined as a time-series process that needs to be first-differenced in order to produce an I (0) process (Wooldridge 2000).

1992, Maddala and Kim 1998). First, the ADF method has a problem of test power (Blough 1992). That is, because the correct form of the ADF model is unknown, different lag lengths cause different results with respect to rejecting the null hypothesis of non-stationarity. Second, the ADF test is unable to detect a structural break in the series, because the implicit assumption of the test is that the deterministic trend is correctly specified (Maddala and Kim 1998). As such, if there is a break in the deterministic trend, then ADF test could lead to a false conclusion that there is a unit root, when in fact there is not (Perron 1989). As a result, to overcome the shortcomings of the usual ADF procedure, this study performs tests for unit roots under structural changes with a known and an unknown break point.

3.2.1. Tests for Structural Change

Before adopting tests for structural changes, this study also introduced the usual ADF procedure to compare between the results without and with structural change. The ADF approach is to test the null hypothesis that a series contains a unit root, or is non-stationary, against the alternative of stationarity (Dickey and Fuller 1979). To illustrate the ADF test, assume that y_t follows $AP(p)$ process,

$$y_t = \varphi^* \, y_{t-1} + \sum_{i=1}^{p-1} \varphi_i \, \Delta y_{t-i} + \alpha + \beta t + u_t \tag{3.1}$$

where u_t is assumed to be white noise. If $\varphi^* = 0$, then y_t contains a unit root. As shown in (3.1), the ADF test involves adding an unknown number of lagged first differences of the dependent variable to capture auto-correlated omitted variables, which otherwise enter the error term u_t. Hence, it is very important to select the appropriate lag-length for the ADF test (Wooldridge 2000).

3.2.1.1. Tests with Known Break Point

Perron (1989) pointed out that if a time-series is stationary around a deterministic time trend due to a permanent shift sometime during the period under consideration, standard tests of unit roots, including ADF, have very low power since they do not consider this change in the slope. In addition, there is a similar loss of power if there has been a shift in the intercept and/or trend. As an alternative approach, therefore, Perron (1989) has suggested a modified ADF test for a unit root with three different models to take into account structural changes. Specifically, if a structural break in time-series is known, then it is relatively simple to adjust the ADF test by including dummy variables that allow a one-time change in the structure occurring at a time T_B, which refers to the time of break. The three different models are parameterized as:

$$\text{Model (A): } y_t = \mu_0 + \delta t + \mu_1 DU_t + u_t$$

where $DU_t = 1$ if $t > T_B$, and 0 otherwise.

$$\text{Model (B): } y_t = \mu + \delta_0 t + \delta_1 DT_t + u_t$$

where $DT_t = 1$ if $t = T_B + 1$, and 0 otherwise.

$$\text{Model (C): } y_t = \mu_0 + \mu_1 DU_t + \delta_0 t + \delta_1 DT_t^* + u_t$$

where $DT_t = t$ if $t > T_B$, and 0 otherwise.

Model (A) considers a one-time change in the intercept of the trend function, which is referred to as the crash model. Model (B) is referred to as the changing growth model to allow a one-time change in the slope of the trend function. Model (C) allows for both changes (slope and intercept) to take place simultaneously. Because the three different models include different dummy variables (DU_t, DT_t, DT_t^*), the null hypothesis of a unit root is different. The null hypothesis is a broken-trend stationary.

3.2.1.2. Tests with Unknown Break Point

Perron's method has been criticized because his procedure is based on a conditional test given a known break point. First, it is implausible that the break points are treated as known a *priori* (Banerjee et al. 1992). Second, because a known break is treated as an exogenous event using a dummy variable, the assumption raises the problem of pre-testing and data-mining with respect to the choice of the break point (Maddala and Kim 1998). As a result, Banerjee et al. (1992) has developed a method that treats the date of the break as unknown a *priori* (BLS test). For this purpose, they employed recursive, rolling and sequential procedures to formulate the asymptotic distributions for the test statistics and tabulate the critical values for these statistics.

Recursive statistics are obtained using sub-samples $t = 1, \ldots, k$ for $k = k_0, \ldots, T$, where k_0 is a start-up value and T is the size of the full sample. That is, the recursive statistics are computed for sub-samples that start with the first sub-sample ranging from 1 to k_0 with $k_0 = 0.25T$, and sequentially increase to cover the full sample. The usual ADF model (equation 3.1) is estimated for each sub-sample, and then the minimum and maximum statistics are chosen and compared to the critical values provided in Table 1 of Banerjee et al. (1992) to test the null of a unit root.

Similarly, rolling statistics are computed using sub-samples that are a constant fraction δ_0 of the full sample, rolling through the sample. The first sub-sample covers 1 to k with $k = 0.3T$, the second sub-sample covers 2 to $k = 0.3T + 1$, and so on. Again, the estimated statistics are compared to the critical values in Table 1 of Banerjee et al. (1992).

The sequential statistics are computed using the full sample and the following adaptation of the ADF model.

$$y_t = \varphi^* y_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta y_{t-i} + \gamma D_t + u_t \tag{3.2}$$

where D_t captures the possibility of a shift or jump in the trend at period t. Banerjee et al. (1992) considered two cases — shift in trend (slope change) and shift in mean value (intercept change):

$$\text{Case A : shift in trend } D_t = \begin{cases} t \text{ if } t > k \\ 0 \text{ if } t \leq k \end{cases}$$

$$\text{Case B: shift in mean } D_t = \begin{cases} 1 \text{ if } \mathrm{t} > k \\ 0 \text{ if } \mathrm{t} \leq \mathrm{k} \end{cases}$$

The unknown date of the break or shift is assumed to occur at k, which is searched out between $0.15T$ and $T - 0.15T$. Minimum statistics for the trend and/or mean-shift models are then compared to the critical values in Table 2 of Banerjee et al. (1992). In addition, F-statistics is used to test the null $H_0 : \varphi^* = \gamma = 0$ in the trend-shift and mean-shift models, and then the maximum values of F-statistics are compared to the critical values in Table 2 of Banerjee et al. (1992).

3.2.2. Test Results for Structural Change

The findings of the ADF test are first reported to compare them with the test results of structural change (Table 3.3). The tests included intercepts and trends, and lag-length was selected based on the Akaike information criteria. The results indicate that for the null hypothesis of a unit root all the data series but U.S. and Canadian processing capacity (PC and CPC) cannot be rejected at the 5% and even 10% significant level.

Table 3.3. Results of ADF unit-root tests

Series	Level	First Difference	Lags
SUP	-2.27	-10.49**	2
EXP	-1.90	-7.17**	2
P	-2.06	-4.83**	2
EX	-0.88	-4.36**	2
HS	-2.24	-6.95**	5
DI	-1.05	-3.49**	4
MR	-0.15	-6.96**	2
W	-2.08	-3.06**	3
CW	-1.26	-3.97**	6
PC	-2.95*	-1.80	5
CPC	-3.19*	-2.28	4
LP	-1.61	-3.96**	4
CLP	-2.20	-3.05**	5

Note:

1. ** and * denote rejection of null hypothesis of a unit root at 5% and 10% significant level.
2. The 5% and 10% critical values for the ADF including a constant and a trend are -2.89 and –2.58.
3. The lag lengths were chosen on the basis of the Akaike information criteria.

3.2.2.1. Test Results with Known Break Point

For Perron's test, descriptive analyses for time-series data were first done to select an appropriate model for each series. Figure 3.1 shows a plot of the logarithm of lumber price series (P), which behaves in correspondence to Model (A). A characteristic of this graph is the conspicuous increase of lumber price series between 1991:1 and 1991:2. Except for this change, the trend appears consistently stable (same slope) over the entire period. The fitted trend (solid line) was estimated from a regression on a constant, a trend and a dummy variable taking a value of 0 prior and at 1991:1, and value 1 afterwards. It is found that four variables have followed the pattern similar to that of Model (A) with different break points. These variables are lumber price (P), U.S. and Canadian processing capacity (PC and CPC), and Canadian log price (CLP).

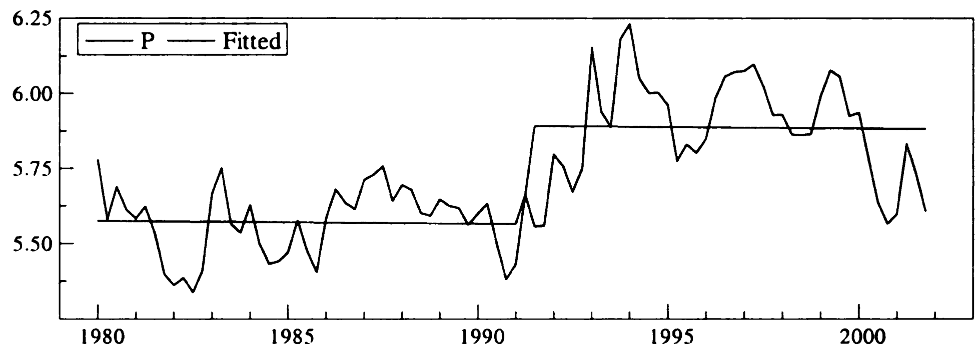


Figure 3.1. A plot of lumber price series

Figure 3.2 shows a plot of the logarithm of U.S. lumber production (SUP), which behaves depicted by Model (C). The break point is in 1991:1, but in this case there appears to be both a sudden change in the intercept of the series in 1991:1 and the slope afterwards. The fitted line is the estimated trend with an intercept dummy (0 prior and at 1991:1, 1 after 1991:1) and a slope dummy (0 prior and at 1991:1 and t after 1991:1). Eight variables follow the pattern depicted by Model (C) with different break points. Among those variables are U.S. lumber production (SUP), housing starts (HS), Canadian lumber export (EXP), exchange rate (EX), disposable income (DI), mortgage rate (MR), wage rate (W), Canadian wage rate (CW), and log price (LP). Note that no variables follow the pattern depicted by Model (B).

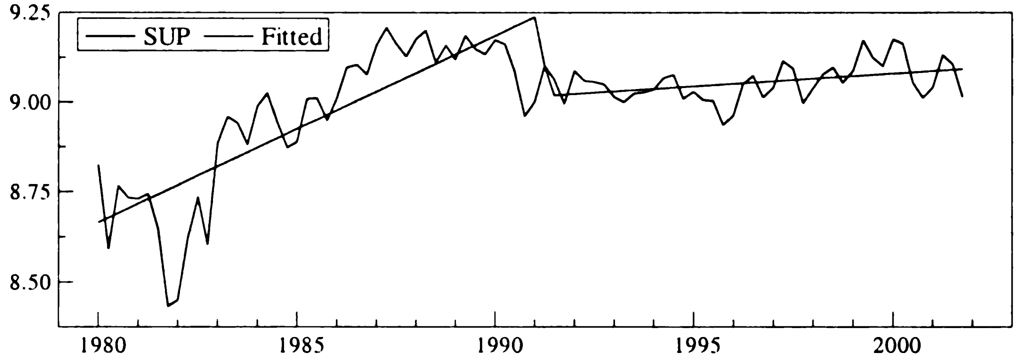


Figure 3.2. A plot of lumber production series

Based on descriptive analyses two unit root tests have been performed. First, the ADF unit root test for split samples was employed in accordance to the break point. The results are summarized in Table 3.4, which indicate that as seen in the ADF test with full samples (Table 3.3), all the data series except for U.S. and Canadian processing capacity (PC and CPC) cannot be rejected with the null hypothesis for a unit root at 5% and 10% significant level.

Then, Perron's test for unit roots was performed with two different models. The results in Table 3.5 are quite consistent with findings without a break point. That is, the unit root hypothesis cannot be rejected for all but two series, or PC and CPC, which behave in correspondence to Model (A). The test statistic was –3.86 for PC, where the critical value at 5% with λ=0.45 (λ is the ratio of pre-break sample size to total sample size) was –3.72~-3.76. While the test statistic of CPC was –3.48, the critical value of 10% with λ=0.45 was –3.44~-3.46. As a result, the null hypothesis of a unit root for both series is rejected at 5% and 10% significance level, respectively.

Table 3.4. Results of ADF unit-root tests with split samples

Series	Period	Observation	Lags	Test statistic
SUP	1980:1-1991:1	45	4	-2.15
	1991:2-2001:4	43	2	-2.22
EXP	1980:1-1991:1	45	3	-0.35
	1991:2-2001:4	43	3	-2.35
P	1980:1-1991:1	45	3	-2.24
	1991:2-2001:4	43	2	-2.26
EX	1980:1-1991:1	45	3	-1.98
	1991:2-2001:4	43	2	-2.39
HS	1980:1-1991:1	45	4	-2.52
	1991:2-2001:4	43	4	-1.83
DI	1980:1-1991:1	45	1	-0.74
	1991:2-2001:4	43	2	-1.90
MR	1980:1-1987:1	29	1	-2.19
	1987:2-2001:4	54	6	-0.72
W	1980:1-1991:1	45	1	-0.25
	1991:2-2001:4	43	2	-2.36
CW	1980:1-1995:1	61	2	-1.46
	1995:2-2001:4	27	2	-1.28
PC	1980:1-1989:4	40	4	-3.16**
	1990:1-2001:4	48	1	-2.66*
CPC	1980:1-1989:4	40	4	-4.47**
	1990:1-2001:4	48	4	-2.95**
LP	1989:1-1991:4	4	-	-
	1991:2-2001:4	48	2	-2.46
CLP	1980:1-1984:4	20	1	-0.94
	1985:1-2001:4	68	1	-1.86

Note:

1. See Table 3.1 for definitions of the variables.
2. ** and * denote rejection of null hypothesis of a unit root at 5% and 10% significant level.
3. The 5% and 10% critical values ($n = 50$) for the ADF tests including a constant and a trend are -2.93 and –2.60.
4. LP for the period of 1989:1-1991:4 was not tested due to small observation.
5. The lag lengths were chosen on the basis of the Akaike information criteria.

Table 3.5. Results of ADF unit-root tests with known break point

Series	Break point	Observation	Model	λ	Lags	Test statistic
SUP	1991:1	88	C	0.51	4	-3.20
EXP	1991:1	88	C	0.51	3	-0.45
P	1991:1	88	A	0.51	2	-2.22
EX	1991:1	88	C	0.51	2	-2.16
HS	1991:1	88	C	0.51	2	-2.83
DI	1991:1	88	C	0.51	4	-2.38
MR	1987:1	88	C	0.33	6	-1.89
W	1991:1	88	C	0.51	4	-2.72
CW	1995:1	88	C	0.69	4	-2.32
PC	1989:4	88	A	0.45	1	-3.86**
CPC	1989:4	88	A	0.45	4	-3.48*
LP	1991:1	52	C	0.17	2	-2.65
CLP	1984:4	88	A	0.22	1	-2.70

Note:

1. See Table 3.1 for definitions of the variables.
2. λ is the ratio of pre-break sample size to total sample size.
3. ** and * denote rejection of null hypothesis of a unit root at 5% and 10% significant level.
4. The 5% and 10% critical value for Model (A) are –3.77 and –3.47 if λ=0.2, -3.76 and -3.46 if λ=0.5. The 5% and 10% critical value for Model (C) are -3.99 and –3.66 if λ=0.2, –4.17 and –3.87 if λ=0.3, -4.24 and –3.96 if λ=0.5, -4.18, and –3.86 if λ=0.7. Critical values are obtained from Table 4B, 5B, and 6B in Perron (1989).
5. According to Perron's procedure, the lag lengths were chosen by a test on the significance of the estimated coefficients for lagged variables in each model.

3.2.2.2. Test Results with Unknown Break Point

The BLS tests were implemented with two lags, and the results are reported in Table 3.6 and 3.7. The recursive and rolling tests in Table 3.6 indicate that all the data series cannot be rejected the unit root null at 5% and 10% significant level. On the other hand, the sequential test in Table 3.7 shows that there is little evidence for rejecting the null hypothesis of a unit root even with allowing the possibility of a break in the series. Only the lumber price series shows weak evidence of a break under a trend-shift regime at the 10% significant level; and by examining the maximum values of F-statistics, it appears that if there was a breakpoint, then it occurred in the first quarter of 1991 (see Figure 3.1).

Table 3.6. Results of recursive and rolling unit-root tests

Series	Observation	ADF statistic	Recursive		Rolling	
			Max t_{DF}	Min t_{DF}	Max t_{DF}	Min t_{DF}
SUP	88	-1.90	-0.70	-2.14	-1.33	-2.46
EXP	88	-2.27	-1.94	-4.29	-0.80	-4.96
P	88	-2.06	-1.18	-2.45	-0.05	-2.36
EX	88	-1.26	0.02	-0.41	-1.48	-2.13
HS	88	-2.08	-1.93	-3.30	-1.85	-3.76
DI	88	-0.88	1.02	-2.49	-1.15	-2.84
MR	88	-2.24	-1.19	-3.87	-1.32	-3.03
W	88	-1.05	-1.64	-3.29	-1.46	-2.68
CW	88	-0.15	-0.72	-2.95	-1.01	-2.87
PC	88	-2.95	0.30	-2.80	-1.26	-2.95
CPC	88	-3.19	-0.04	-2.70	-0.65	-3.61
LP	52	-1.61	-0.52	-2.65	-1.07	-3.60
CLP	88	-2.20	-0.57	-2.18	-1.30	-3.03

Note:

1. ** and * denote rejection of null hypothesis of a unit root at 5% and 10% significant level; The 5% and 10% critical values for the ADF including a constant and a trend are -2.89 and –2.58 if n =88, -2.93 and –2.60 if n =52; The lag lengths (k =2) was chosen on the basis of the Akaike information criteria.
2. The 5% and 10% critical values for Min t_{DF} (sample size T =100) are –4.33 and –4.00 if recursive tests, -5.01 and –4.71 if rolling tests. The 5% and 10% critical values for Max t_{DF} (sample size T =100) are –1.99 and –1.73 if recursive tests, -1.49 and –1.31 if rolling tests. Critical values are from Table 1 in Banerjee et al. (1992).

Table 3.7. Results of sequential unit-root tests

Series	Observation	Mean-shift statistic		Trend-shift statistic	
		Min t_{DF}	Max F	Min t_{DF}	Max F
SUP	88	-3.16	12.33	-3.65	11.43
EXP	88	-3.76	13.81	-3.58	13.33
P	88	-3.44	13.64	-4.21*	14.30*
EX	88	-2.01	6.52	-2.91	7.46
HS	88	-2.52	5.21	-2.57	5.30
DI	88	-1.82	5.24	-1.72	5.57
MR	88	-3.12	7.46	-1.99	6.99
W	88	-1.74	8.80	-2.17	8.40
CW	88	-1.06	8.09	-1.70	9.72
PC	88	-2.12	7.39	-1.78	8.29
CPC	88	-3.17	5.07	-2.25	4.61
LP	52	-1.18	0.74	-1.36	3.93
CLP	88	-2.81	4.38	-3.43	5.90

Note:

1. ** and * denote rejection of null hypothesis of a unit root at 5% and 10% significant level; The 5% and 10% critical values for the ADF including a constant and a trend are -2.89 and –2.58 if n =88, -2.93 and –2.60 if n =52; The lag lengths (k =2) was chosen on the basis of the Akaike information criteria.
2. The 5% and 10% critical values for Min t_{DF} (sample size T =100) are –4.48 and –4.20 if trend-shift tests, -4.80 and –4.54 if mean-shift tests. The 5% and 10% critical values for Max F (sample size T =100) are 16.30 and 13.64 if trend-shift tests, 18.62 and 16.20 if mean-shift tests. Critical values are from Table 2 in Banerjee et al. (1992).

3.3. Summary and Conclusions

This study collected quarterly data for 1980-2001. For the U.S., the data include lumber production and prices, log prices, housing starts, disposable income, lumber processing capacity, sawmilling wage rates, and mortgage rates. For Canada, the data include lumber exports, log prices, bilateral exchange rates, lumber processing capacity, and sawmilling wage rates. Because U.S. log prices in the PNW can only be traced back to the first quarter of 1989, this study compiled two datasets, with dataset I containing every variables but PNW log prices for 1980:1-2001:4 and dataset II all the variables for 1989:1-2001:4.

This study employed the usual ADF and structural change procedures to test for the existence of unit roots in the time-series. The usual ADF test shows that there is no evidence for rejecting the unit root null hypothesis for all the series but *CP* and *CPC* series. The Perron's tests with known break points confirm the findings of the usual ADF test. Furthermore, the BLS tests with unknown break points show little evidence of a break except for lumber price series. Based on these results, therefore, it is safe to conclude that *CP* and *CPC* series cannot be treated as non-stationary for further time-series analyses.[13]

In contrast, this study found that the lumber price series to be non-stationary in view of the following facts. First, Perron's test for the price series indicates that there is little evidence of a structural change around the first quarter of 1991. In addition, with the BLS tests, the unit root null was rejected on the borderline of the 10% significant level.

[13] Note that this study employed interpolation of the annual data for processing capacities for the U.S. and Canada to obtain quarterly observations. In conjunction with a strong trend in the capacity series, this treatment may influence the results of unit root tests and their usefulness in the estimation.

Second, the structural change tends to under-reject the null of no co-integration (Gregory et al. 1996). In other words, the structural change could change in the number of co-integrating vector (Quintos 1995, Maddala and Kim 1998). However, the co-integration rank test in this study is shown to be sustainable regardless of whether or not a structural break exists in the series. For example, with the first quarter of 1991 as the break point the full sample (1980:1-2001:4) was separated into two different periods — Sample I (1980:1-1990:4) and Sample II (1991:1-2001:4). The Johansen co-integration test was then applied three endogenous variables, including EXP, SUP, and P, in order to find a change in the rank of co-integration space.[14] As shown in Table 3.8, this study found that with all three cases there is only one co-integration equation at 5% significant level, or rank one ($r = 1$). This implies that the structural change in the lumber price series does not change the number of co-integration rank. In combination, therefore, it is acceptable that the lumber price series can be treated as non-stationary and thus included in further analysis.

Table 3.8. Results of Johansen co-integration test with structural change

Sample	Full sample (1980:1-2001:4)		Sample I (1980:1-1990:4)		Sample II (1991:1-2001:4)		5% Critical value
Null hypothesis	E-Value	Trace statistics	E-Value	Trace statistics	E-Value	Trace statistics	
R=0	0.370	42.34*	0.560	40.33*	0.426	36.41*	29.68
R≤1	0.031	3.54	0.145	7.48	0.222	11.96	15.41
R≤2	0.011	0.91	0.029	1.22	0.020	0.87	3.76

Note:
1. * denotes rejection of the null hypothesis at the 5% level.
2. With four lags, the tests indicate 1 co-integration equation at 5% level ($r = 1$).

[14] Given small observations for sample I and II ($n = 44$ for each), the Johansen co-integration test with structural change was only applied to the three endogenous variables (see Chapter 5.2 for detailed co-integration tests).

CHAPTER 4

A STRUCTURAL MODEL FOR U.S. SOFTWOOD LUMBER MARKET

This chapter provides empirical analyses for the structural models, a simultaneous system of equations (SSE) and a reduced-form equation (RFE), for the U.S. softwood lumber market. First, an SSE model is used to characterize the relationships among all the variables associated with the lumber market. The main tasks of the SSE model are to assess the price and volume impacts of a restrictive regime and to derive the price elasticities used for welfare analysis later. Second, several previous studies have used an RFE to assess the effects of the U.S. trade restrictions on lumber prices and Canadian share in the U.S. market. Given that, estimating an RFE for lumber prices and Canadian export supply in this study makes it possible to detect causes and effects of the potential variations between this and other studies. This step also helps discover how and why results derived may differ from those obtained from the SSE and VAR models in this study. This chapter is organized as follows. First, the theoretical framework for the structural models is discussed. Next, economic models as well as the estimation procedures for both the SSE and RFE models are presented. Then, the empirical results are discussed. Finally, this chapter ends with summary and comparison of the results with other studies.

4.1. Theoretical Framework

Assuming that softwood lumber is a homogenous commodity, this study develops models for the U.S. softwood lumber market in aggregation. This aggregate analysis serves as a useful approach to assessing the overall impacts of the U.S. trade restrictions — the 1986 MOU and the 1996 SLA (Wear and Lee 1993). However, this approach does not take into account the spatial structure of regional markets and transportation costs. Hence, consumers are assumed to be indifferent to lumber sources, and U.S. import demand and Canadian export supply can be seen as a function of a single representative price.

Figure 4.1 illustrates the approach for estimating market impacts of a trade restriction. Specifically, Figure 4.1 shows demand and supply in the U.S.-Canada softwood lumber trade with U.S. import demand and Canadian export supply. In this model, the action of a trade restriction, such as an export tax and a quota, on softwood lumber imported from Canada can be envisioned as an upward shift in Canadian export supply. As a result, the new equilibrium is achieved at a reduction in imported lumber quantity from Canada ($Q_0 \rightarrow Q'$), an increase in U.S. prices of both imported and domestic lumber ($P_0 \rightarrow P'_{US}$), and a decrease in the price received by Canadian producers ($P_0 \rightarrow P'_{CA}$).

The extent of those changes in the U.S. and Canada mainly depends on the price elasticities of Canadian export supply and U.S. import demand. For example, if Canadian export supply is elastic (sensitive to price changes), price changes by a trade restriction will have a great effect on the quantity of lumber imported. In other words, if market price were to change even slightly, the quantity imported would change significantly. Similarly, if U.S. import demand is elastic, a restriction will cause smaller price changes and larger import quantity changes in the U.S. The price elasticity of U.S. import demand depends on the price responsiveness of U.S. demand and of U.S. production for lumber (Adams 2003).

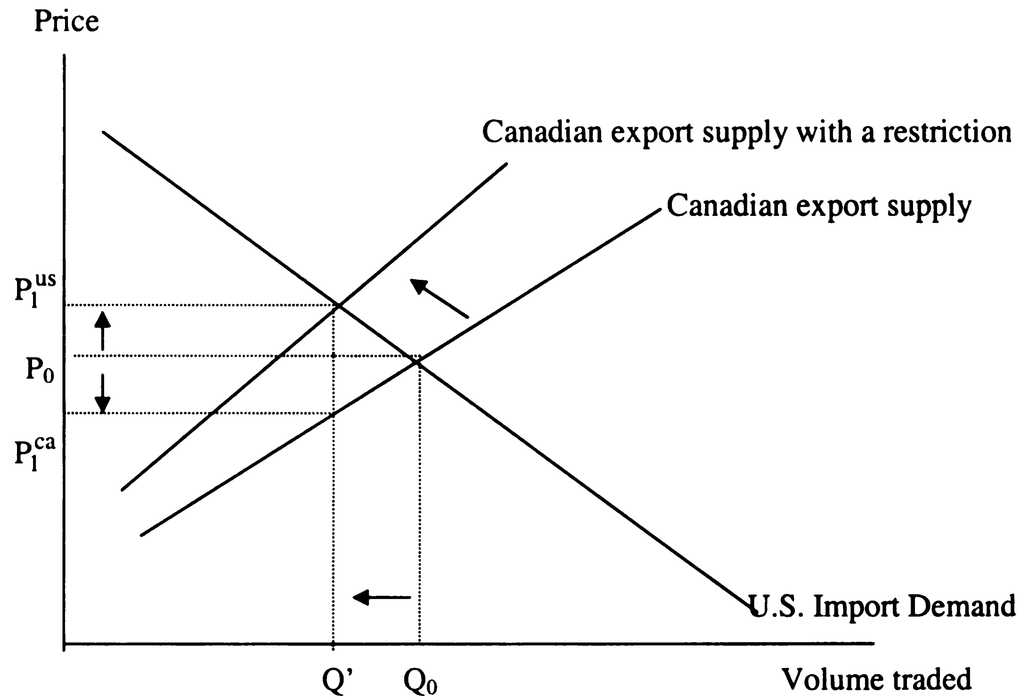


Figure 4.1. Structure of the U.S. softwood lumber market.

4.1.1. Demand for Softwood Lumber

The derived demand for softwood lumber can be obtained from the total cost function of the primary end uses, which is known as the duality between production and cost function (Mas-Colell et al. 1995). Specifically, according to economic theory, the determinants of the cost of end uses for softwood lumber are price, the level of end outputs, and the costs of all other factors. As a result, the total cost function (TC) can be specified in the following;

$$TC = f(P, Y, C)$$

where P is lumber price, Y is the level of end outputs, and C is the costs of all other factors such as the costs of labor and capital, and all other materials. The derived demand for lumber (DEM) then can be obtained directly from the total cost function by using Shephard's Lemma (Mas-Colell et al. 1995).

$$DEM = \frac{\partial TC}{\partial p} = f(Y, P, C)$$

The demand for end outputs is mainly derived from the demand for housing, including residential construction and repair & remodeling. The demand for housing is a function of housing starts (HS), disposable income (DI), and mortgage rates (MR). Housing demand will rise with an increase in housing starts and disposable income. Additionally, the demand for housing will rise with a decrease in mortgage rates. Therefore, by substituting housing starts, disposable income, and mortgage rate for the end outputs, the derived demand for lumber in the U.S. can be specified in the following;

$$DEM = f(HS, DI, MR, P)$$

4.1.2. Supply of Softwood Lumber

The supply of softwood lumber can be derived from the profit function, which is known as Hotelling's lemma (Mas-Colell et al. 1995). Specifically, individual sawmills can be viewed as possessing a profit function (P, X), where P is lumber price and X is the price (cost) of all other factors. Then, the supply of lumber can be obtained directly from the profit function by using Hotelling's lemma (Mas-Colell et al. 1995).

$$SUP = \frac{\partial \pi}{\partial P} = f(P, X)$$

The cost of all other factors is a function of the prices of input factors such as labor (W), capital (V), logs (LP), and the price of all other materials (M). The supply of softwood lumber will fall with a rise in prices of input factors. Note that this study used lumber processing capacity (PC) as a proxy for capital input (V) in order to reflect embodied technical change in lumber production. Hence, the supply of lumber will rise with an increase in lumber processing capacity. Therefore, by substituting wage, lumber capacity and log prices for the price of all other factors, the supply of lumber in the U.S. can be specified in the following;

$$SUP = f(W, PC, LP, P)$$

4.1.3. Canadian Export Supply of Softwood Lumber

Canadian export supply of softwood lumber also can be derived from the profit function by using Hotelling's lemma (Mas-Colell et al. 1995). As such, Canadian export supply of lumber can be specified as a function of the prices of input factors such as labor (CW) and log (CLP), lumber processing capital (CPC), and exchange rates (EX). Note that Canada's export supply of lumber should rise with a depreciation of Canadian dollar against U.S. dollar via decreasing Canadian real costs relative to those in the U.S. Hence, exchange rate is an important factor determining the relative cost competitiveness between the U.S. and Canada. As a result, Canadian export supply of lumber can be specified in the following;

$$EXP = f(CW, CPC, CLP, EX, P)$$

4.2. Econometric Models and Estimation Methods

4.2.1. A Simultaneous System of Equations (SSE) Model

In order to develop and estimate an SSE model for the U.S. softwood lumber market, a four-equation system is postulated based on the discussion in the previous section. Specifically, the U.S. demand (DEM) is specified as a function of lumber prices (P), housing starts (HS), personal disposable income (DI), and mortgage rates (MR). Similarly, the U.S. supply (SUP) is specified as a function of lumber prices (P), log prices (LP), processing capacity (PC), sawmilling wage rates (W), trend (T), and shocks from the log market — DUM_1 for the federal harvest reductions in the PNW and DUM_2 for the Asian financial crisis. Notice that these two dummies are only added when estimating with dataset I, which does not have the data for U.S. log prices.[15] Canadian export supply of softwood lumber (EXP) is specified as a function of lumber price (P), Canadian log prices (CLP), Canadian processing capacity (CPC), sawmilling wage rates (CW), exchange rates (EX), and dummies for the trade restrictions (MOU and SLA). These three equations include seasonality (for quarters 2-4). Finally, as an identity in quantities, the domestic demand for lumber is defined as the sum of domestic supply and Canadian export supply less U.S. exports ($LGEX$). Therefore, the SSE model is expressed as the following:

[15] Recent studies show that the log market dynamics have influenced the lumber market. Yin (2001) found that at least 70% of the log price changes could have been caused by the federal harvest reductions, which in turn contributed to higher and more volatile prices for lumber in the early 1990s. In addition, Yin and Xu (2002) found that export log prices lead the formation of log prices in the PNW, which in turn affects the domestic lumber prices. Thus, it is essential to incorporate the dummy variables to capture effects of the federal harvest reductions and Asian financial crisis.

$$DEM_t = \alpha_1 + \alpha_2 \mathrm{P_t} + \alpha_3 \mathrm{HS_t} + \alpha_4 \mathrm{DI_t} + \alpha_5 \mathrm{MR_t} + \mathrm{SEA} + u_t$$
$$SUP_t = \beta_1 + \beta_2 \mathrm{P_t} + \beta_3 \mathrm{PC_t} + \beta_4 \mathrm{W_t} + \beta_5 \mathrm{LP_t} + \beta_6 \mathrm{T} + \beta_7 \mathrm{DUM_1} + \beta_8 \mathrm{DUM_2} + \mathrm{SEA} + u_t$$
$$EXP_t = \gamma_1 + \gamma_2 \mathrm{P_t} + \gamma_3 \mathrm{CPC_t} + \gamma_4 \mathrm{CW_t} + \gamma_5 \mathrm{CLP_t} + \gamma_6 \mathrm{EX_t} + \gamma_7 \mathrm{MOU} + \gamma_8 \mathrm{SLA} + \mathrm{SEA} + u_t$$
$$DEM \equiv SUP + EXP - LGEX$$

where α's, β's, and γ's are coefficients to be estimated, t denotes the time period of the observation, and ε's are i.i.d error terms. With an identity to close the system, this study has four equations in four endogenous variables (DEM, SUP, EXP and P), which is similar to the approach of Myneni et al. (1994). Unlike that work, however, this study has not only a full coverage of the two major episodes of the trade restrictions, but also a substantially expanded list of variables.[16] Furthermore, the use of quarterly instead of annual data results in more observations for analysis.

The trend variable in the domestic equation is taken to reflect disembodied technical change in lumber production. The processing capacity is a proxy for capital input (including embodied technical change), but only annual data are available (Howard 2000). Hence, the annual data is interpolated to obtain quarterly observations in a way similar to Zhang (2001). Compared to Sarker (1996) that only considered demand factors, Canadian log prices, lumber processing capacity, and sawmilling wage rates are included in the Canadian export equation. The two dummy variables (MOU and SLA) are used to allow for an intercept shift in Canadian export supply during the periods of the trade restrictions. As in Myneni et al. (1994), the SSE model assumes that a significant drop in Canadian export supply due to a trade restriction will lead to a significant price increase and thus domestic production increase.

[16] Previous studies on the impacts of the trade restrictions have incompletely covered the 1986 MOU and the 1996 SLA.

Notice that the SSE model satisfies the order condition for identification, which requires that the number of exogenous variables excluded from any equation must be as least as large as the number of endogenous variables in that equation (Wooldridge 2000). In the demand equation (DEM), there are two endogenous variables (DEM and P) and more than two excluded exogenous variables such as LP , PC , and W . Similarly, the condition is satisfied for the supply and Canadian export supply equation. It should be noted, however, that the requirement is the only necessary condition for identification. Hence, this study assumes that an equation that satisfies the order condition is identified for consistent estimation (Wooldridge 2000).

The procedure of three-stage least squares (3SLS) is employed to estimate the SSE model. This method constitutes both a seemingly unrelated regression (SUR) approach and an instrumental variable (IV) approach (Wooldridge 2000). Specifically, an SUR approach is employed to use the information from the full system of equations to obtain efficiency. An IV approach is used to generate consistent estimators. This study uses all the exogenous variables in the three equations as instrumental variables (IV). As a result, a 3SLS is called a full information system method. For that reason, 3SLS produces more efficient estimators than those of limited information methods, such as two-stage least squares (2SLS) and limited information maximum likelihood, which estimate one equation's parameters only using information in that equation (Darnell 1994).

4.2.2. A Reduced-Form Equation (RFE) Model

An RFE is a reduced version of an SSE model. It is defined as a linear equation in which an endogenous variable is a function of exogenous variables and unobserved errors (Wooldridge 2000). To understand this concept, consider a two-equation structural model as follows:

$$\begin{aligned} y_1 &= \alpha_1 y_2 + \beta_1 z_1 + u_1 \\ y_2 &= \alpha_2 y_1 + \beta_2 z_2 + u_2 \end{aligned} \tag{4.1}$$

where z_1 and z_2 are exogenous variables, so that they are not correlated with error terms, u_1 and u_2. Assuming that we focus on estimating the first equation, it is possible to express the two equations as the reduced form for y_2. That is, if we solve the two equations for y_2,

$$y_2 = \Pi_{21} z_1 + \Pi_{22} z_2 + v_2 \tag{4.2}$$

where $\Pi_{21} = \frac{\alpha_2 \beta_1}{(1-\alpha_2\alpha_1)}$, $\Pi_{22} = \frac{\beta_2}{(1-\alpha_2\alpha_1)}$, and $v_2 = \frac{(\alpha_2 u_1 + u_2)}{(1-\alpha_2\alpha_1)}$. Equation (4.2) can be estimated by ordinary least squares (OLS) regression of y_2 on z.

Given the nature of an RFE, exogenous variables contained in the SSE model can be expressed as the determinants of endogenous variables such as lumber prices (P) and Canadian export supply (EXP). The determinants of the lumber demand include housing starts (HS), disposable income (DI), and mortgage rates (MR). The determinants of lumber supply include log prices (LP), lumber processing capacity (PC), sawmilling wage rates (W), and shocks from the log market (DUM_1 and DUM_2). The determinants of Canadian export supply include Canadian log prices (CLP), Canadian lumber processing capacity (CPC), Canadian sawmilling wage rates (CW), exchange rates

(EX), and dummies for the trade restrictions (MOU and SLA). The RFE model can be specified as:

$$P = f(HS, DI, MR, LP, PC, W, CLP, CPC, CW, EX, DUM_1, DUM_2, MOU, SLA)$$

or,

$$EXP = f(HS, DI, MR, LP, PC, W, CLP, CPC, CW, EX, DUM_1, DUM_2, MOU, SLA)$$

Note that DUM_1 and DUM_2 are only used when estimating dataset I, where U.S. log prices are unavailable. In addition, dummy variables for trade restrictions are used in different ways. As in the SSE, a single dummy was first introduced to cover all the years under a specific policy — MOU during 1986:4-1991:3 and SLA during 1996:2-2001:1. Another way was to use multiple dummies for the individual years under the SLA (Zhang 2001) to look into the potential ramifications of these alternative specifications. Finally, whether or not each of the variables in the above specification is included in the final model will be determined by model performance.

For comparison, among studies of the lumber price impact of the trade restrictions, Lindsey et al. (2000) included housing starts growth rates for the current and previous year, previous year's price for lumber, Japanese GDP. The effects of the trade restrictions were estimated in two ways. One way was to use a single dummy including both the MOU and the SLA between 1986-1998. The other way was to make distinction between the SLA during 1996-1998 and the earlier forms of restriction during 1986-1995. Zhang's (2001) final model had previous year's price for lumber, housing starts, lumber production capacity, sawmilling wage rates, lumber productivity, and exchange rates. He also included multiple dummies for the individual years of the SLA during 1996:2-2000:1 and a single dummy for the MOU during 1986:4-1991:3 to estimate the trade restriction effects. On the other hand, Wear and Lee (1993) included housing starts, GNP, time trend, U.S. lumber and log exports, Canadian lumber consumption, exchange rates, and a single dummy for the MOU (1986-1990) in assessing the impact of the trade restrictions on Canadian market share

For estimating the RFE, data series were detrended because of better model performance in terms of goodness-of-fit and coefficient significance. Detrending data were obtained from the residuals derived from the regression of each of dependent and independent variables on a constant, and time trend (Wooldridge 2000).

4.3. Empirical Results

4.3.1. SSE Model

This study adopted both logarithmic and linear functional forms to estimate the SSE model with dataset I (1980:1-2001:4) and dataset II (1989:1-2001:4). Linear function forms showed better model performance, in terms of a better fit and the number of significant variables. This presentation thus focuses on the results derived from the linear regression, which are reported in Table 4.1. Overall, the SSE model fits the data well. The R^2 is quite high, above 0.8. In addition, the Durbin-Watson (DW) indicates that the hypothesis of no serial correlation cannot be rejected at the 1% level. More importantly, the coefficients have the expected/acceptable signs, and most of them are significant. To mitigate the influence of collinearity, the demand equation was estimated with either mortgage rates or personal disposable income for the two datasets, depending on which alternative would give rise to a better fit.[17]

[17] This study detected a high correlation between mortgage rates and personal disposable income (–0.93).

4.3.1.1. Results of Dataset I

All the coefficients of lumber demand are significant at the 5% level and have the expected signs. Specifically, demand for softwood lumber is negatively related to lumber prices and mortgage rates, but positively related to housing starts. All coefficients but DUM_2 in supply equation are statistically significant and have expected signs. That is, domestic supply of softwood lumber is positively related to lumber prices, production capacity, and wage rates. Further, domestic supply of softwood lumber is negatively related to time trend and DUM_1. The negative time trend indicates that the slow down of the disembodied technical change during the 1980-2001 period, but the coefficients imply only a very small effect (less than 1%). In addition, the coefficient of the dummy for the log market shock, resulting from the spotted owl debate and federal harvest reductions, indicates a significant negative effect on U.S. production, -5.4%. However, the dummy for the Asian financial crisis (DUM_2) is insignificant.

Most coefficients in Canadian export equation are statistically significant and have the expected signs. In particular, Canadian export supply is positively related to lumber prices, exchange rates, and Canadian production capacity. On the other hand, Canadian export supply is negatively related to the dummy variable for the MOU. On a percentage basis, the MOU depressed Canadian exports by about 8%. However, the dummy for the SLA is insignificant, implying that it had little impact on Canadian exports.

To better gauge the influences of explanatory variables on demand, and domestic and export supply, this study calculated the elasticities (Table 4.2) corresponding to coefficient estimates in Table 4.1 and mean values of the variables in Table 3.2. Overall, the prices elasticities are very low. Specifically, the price elasticity of lumber demand is relatively very low (-0.13), compared to other market variables. For example, when lumber prices increase by 1%, demand for softwood lumber falls by 0.13%, holding other factors fixed. On the other hand, when housing starts increase by 1%, demand for lumber increases by 0.33%. In addition, demand for lumber decreases by approximately 0.20%, as mortgage rates go up by 1%. Additionally, the impact of lumber prices on domestic production is smaller than the influences of other variables. For instance, the elasticity of lumber price is 0.26, but the elasticities of wage rates and production capacity are 1.08 and 1.88, respectively. Finally, the price elasticity of Canadian export supply is extremely low (0.13). In other words, when lumber price increases by 1%, Canadian lumber exported to the U.S. only increases by 0.13%, holding other factors fixed. In contrast, the high price elasticity of exchange rates shows that when the U.S. dollar depreciates against the Canadian dollar by 1%, softwood lumber imports from Canada increase by almost the same proportion (0.91).

4.3.1.2. Results of Dataset II

Despite a small sample size ($n=52$), the SSE model with dataset II provides useful information on the U.S. lumber market in the 1990s. Because of the high correlation between mortgage rate and personal disposable income (-0.93) with dataset II, lumber demand is specified as a function of lumber price, housing starts, personal disposable income, and seasonality (for quarters 2-4), which gives a better goodness-of-fit. The results showed quite high R^2 (about 0.88), and all the coefficients are significant at the 5% level and have the expected signs. Demand for softwood lumber is negatively related to lumber prices, but positively related to housing starts and disposable income. Notice that the elasticity of lumber price (-0.08) with dataset II is lower than that from dataset I (-0.13, Table 4.2). It implies that the price elasticity for lumber demand has become lower in the 1990s. The results show that housing starts (0.34) and disposable income (0.46) have significant impacts on demand for softwood lumber in the 1990s.

Note that, instead of the dummies for log market shocks, log prices were directly used with dataset II. All coefficients but production capacity in supply equation are significant at the 5% level and have the expected signs. Domestic supply of softwood lumber is positively associated with lumber prices and wage rates, but negatively related to log prices. It is now proven that log prices have a significant negative effect on lumber supply. In addition, the elasticity of lumber prices (0.15) is slightly lower than that from dataset I (0.26, Table 4.2).

Canadian export supply equation has only two significant variables — lumber price and exchange rate, with the expected signs. Canadian export supply is positively related to lumber prices and exchange rates. The elasticity of lumber prices is lower (0.15). The elasticity of exchange rates is relatively high (0.62), but lower than those from dataset I (0.91~0.92, Table 4.2). However, the dummy variables for the 1986 MOU and the 1996 SLA are not statistically significant. Notice that the small sample size for the 1986 MOU (11-quarter period) with dataset II may influence the significance of the dummy.

Table 4.1. Results of the SSE model for the U.S. lumber market

Equation	Explanatory variables	Dataset I (1980:1-2001:4)		Dataset II (1989:1-2001:4)	
		Coefficient	*t*-statistic	Coefficient	*t*-statistic
DEM	Intercept	12658.35	19.49*	3983.06	5.40*
	Lumber price	-5.07	-3.67*	-2.74	-2.15*
	Housing starts	10.94	10.92*	11.91	4.35*
	Mortgage rate	-238.94	-17.96*		
	Disposable income			0.86	4.33*
	Seasonal (2nd quarter)	-628.93	-3.11*	-479.83	-1.61
	Seasonal (3rd quarter)	-680.19	-3.56*	-604.57	-2.18*
	Seasonal (4th quarter)	-781.51	-4.61*	-650.12	-3.57*
	R^2	0.89		0.87	
	DW	2.05		1.72	
SUP	Intercept	-53194.37	-17.24*	-12233.04	-4.90*
	Lumber price	7.09	7.52*	3.88	3.23*
	Wage rate	941.24	6.44 *	1940.87	8.70*
	U.S. Capacity	1.40	19.27*	1.10	1.15
	Log price			-2.01	-2.94*
	Trend	-170.48	-16.93*	13.58	4.06*
	DUM_1 (1991:1-1993:4)	-465.18	2.51*		
	DUM_2 (1997:3-1998:4)	259.38	1.19		
	Seasonal (2nd quarter)	471.82	4.31*	459.41	4.55*
	Seasonal (3rd quarter)	-147.53	-1.08	192.51	1.86
	Seasonal (4th quarter)	-308.71	-2.58*	-27.96	-0.27
	R^2	0.89		0.80	
	DW	1.86		1.60	
EXP	Intercept	-5547.64	-11.37*	-13382.70	-3.11*
	Lumber price	1.64	2.37*	1.76	3.05*
	Exchange rate	2490.34	7.05*	1845.23	3.65*
	Canadian Capacity	0.20	9.29*	0.09	0.66
	Canadian Log price	-0.43	-0.18	10.84	1.22
	Canadian Wage rate	-6.64	-0.15	8.32	0.25
	MOU (1986:4-1991:3)	-288.56	-3.55*	-130.24	-1.47
	SLA (1996:2-2001:1)	-139.12	-1.50	-98.52	-0.76
	Seasonal (2nd quarter)	523.22	6.56*	410.30	5.45*
	Seasonal (3rd quarter)	444.47	5.46*	241.63	2.66*
	Seasonal (4th quarter)	-21.46	-0.28	-2.59	-0.03
	R^2	0.88		0.93	
	DW	1.41		2.00	

Note:

1. Numbers in parentheses are *t*-values; * indicates significance at the 5% level.
2. The Durbin-Watson test indicates that the hypothesis of no serial correlation cannot be rejected at the 1% level in all but one case. Even for the Canadian export equation with dataset I and for the supply equation with dataset II, the statistics are well within the 5% fail-to-reject region.

Table 4.2. Elasticities for the U.S. softwood lumber market

Equation	Variables	Dataset I	Dataset II
DEM	Lumber price	-0.13	-0.08
	Housing starts	0.33	0.34
	Mortgage rates	-0.21	
	Disposable income		0.46
SUP	Lumber price	0.26	0.15
	Wage rates	1.08	2.41
	Processing capacity	1.88	
	Log prices		-0.14
	Trend	-0.93	0.04
EXP	Lumber price	0.13	0.15
	Exchange rates	0.91	0.62
	Processing capacity (CA)	0.37	0.16

Note:

1. Data means from Table 3.2 and coefficient estimates from Table 4.1 are used to derive the elasticites.

4.3.2. The RFE Model

The OLS was used to estimate the RFE models for lumber price and Canadian export supply with dataset I (1980:1-2001:4) and dataset II (1989:1-2001:4). A logarithmic functional form was adopted with detrended data series. Results of the RFE estimation are reported in Tables 4.3 and 4.4. Overall, the RFE models fit the data well (R^2 is above 0.7). In addition, $AR(q)$ serial correlation test (F test) and Breusch-Godfrey test (LM test) indicate that the hypothesis of no serial correlation cannot be rejected at the 5% level. Breusch-Pagan ($B-P$) test indicates that the hypothesis of no heteroskedasticity cannot be rejected at the 5% level.

Table 4.3. Results of reduced-form equations for lumber price

Variables	Dataset I (1980:1-2001:4)		Dataset II (1989:1-2001:4)	
Housing starts	0.45 (3.56)*	0.47 (3.64)*	0.49 (1.97)*	0.66 (2.71)*
Mortgage rate	-0.11 (-0.44)	0.13 (0.50)	0.01 (0.01)	0.11 (0.40)
Wage rate	0.29 (0.30)	2.18 (2.32)*	-0.89 (-1.45)	-1.05 (-1.69)
Capacity	-3.52 (-1.91)	-4.29 (-2.31)*	-3.39 (-0.77)	-4.58 (-1.34)
Log price			0.47 (2.66)*	0.60 (3.86)*
Canadian log price	0.67 (3.87)*	0.84 (5.74)*		
Canadian wage rate	-0.03 (-0.10)	0.26 (0.79)		
Exchange rate	-0.06 (-0.14)	-0.74 (-2.12)*	0.37 (0.53)	-0.12 (-0.18)
MOU_{87-91} (1986:4-1991:3)	0.11 (1.99)*	0.13 (2.15)*	-0.04 (-0.76)	0.003 (0.07)
SLA_{96-97} (1996:2-1997:1)	0.07 (1.18)		0.09 (1.43)	
SLA_{97-98} (1997:2-1998:1)	0.11 (1.27)		0.04 (0.64)	
SLA_{98-99} (1998:2-1999:1)	0.08 (0.91)		-0.03 (-0.43)	
SLA_{99-00} (1999:2-2000:1)	0.19 (3.35)*		0.09 (1.28)	
SLA_{00-01} (2000:2-2001:1)	-0.11 (-1.77)		-0.13 (-1.95)	
SLA_{96-01} (1996:2-2001:1)		0.05 (1.47)		0.02 (0.47)
DUM_1 (1991:1-1993:4)	0.15 (2.17)*	0.13 (1.85)		
DUM_2 (1997:3-1998:4)	-0.11 (-1.23)	-0.08 (-1.44)		
SEA_1 (2nd quarter)	-0.14 (-2.70)*	-0.15 (-2.70)*	-0.18 (-2.17)*	-0.24 (-2.84)*
SEA_2 (3rd quarter)	-0.17 (-3.67)*	-0.18 (-3.29)*	-0.19 (-2.72)*	-0.24 (-3.29)*
SEA_3 (4th quarter)	-0.12 (-3.39)*	-0.11 (-2.87)*	-0.14 (-2.95)*	-0.14 (-3.05)*
Intercept	0.05 (1.55)	0.07 (1.83)	0.13 (2.30)*	0.15 (2.48)*
R^2	0.71	0.67	0.82	0.76

Note:

1. Numbers in parentheses are *t*-values, and * indicates significance at the 5% levels.
2. The hypothesis of no serial correlation cannot be rejected at the 5% level.
3. Variables are in logarithmic terms with detrending. To mitigate collinearity between U.S. and Canadian capacity in both datasets (0.94), the equations were estimated with either U.S. or Canadian capacity, and the estimation that gives a better fit is reported. To mitigate collinearity between U.S. and Canadian wage rates in dataset II (-0.74), the equations were estimated with either U.S. or Canadian wage rates and reported the estimation that gives a better fit.

Table 4.4. Results of reduced-form equations for Canadian export supply

Variables	Dataset I (1980:1-2001:4)		Dataset II (1989:1-2001:4)	
Housing starts	0.62 (8.09)*	0.63 (8.39)*	0.45 (4.08)*	0.49 (4.78)*
Mortgage rate	0.04 (0.36)	0.06 (0.54)	-0.09 (-0.81)	-0.12 (-1.04)
Wage rate	-1.42 (-3.37)*	-1.45 (-3.73)*	-0.51 (-1.92)	-0.51 (-1.88)
Capacity			2.90 (1.53)	2.59 (1.57)
Canadian capacity	1.62 (4.68)*	1.60 (4.75)*		
Log price			0.14 (1.98)*	0.14 (2.28)*
Canadian log price	0.07 (0.68)	0.09 (0.98)		
Canadian wage rate	0.35 (1.77)	0.39 (2.18)*		
Exchange rate	0.62 (2.79)*	0.59 (2.68)*	0.17 (0.54)	0.05 (0.17)
MOU_{87-91} (1986:4-1991:3)	-0.13 (-3.31)*	-0.13 (-3.42)*	-0.06 (-2.53)*	-0.02 (-0.93)
SLA_{96-97} (1996:2-1997:1)	-0.02 (-0.46)		0.04 (1.40)	
SLA_{97-98} (1997:2-1998:1)	-0.06 (-1.21)		0.01 (0.18)	
SLA_{98-99} (1998:2-1999:1)	-0.10 (-1.89)		-0.04 (-1.23)	
SLA_{99-00} (1999:2-2000:1)	-0.01 (-0.13)		-0.04 (-1.31)	
SLA_{00-01} (2000:2-2001:1)	-0.05 (-1.37)		-0.03 (-1.08)	
SLA_{96-01} (1996:2-2001:1)		-0.03 (-1.63)		-0.01 (-0.59)
DUM_1 (1991:1-1993:4)	-0.01 (-0.33)	-0.02 (-0.46)		
DUM_2 (1997:3-1998:4)	0.06 (1.12)	0.01 (0.30)		
SEA_1 (2nd quarter)	-0.10 (-2.99)*	-0.10 (-3.15)*	-0.05 (-1.34)	-0.06 (-1.68)
SEA_2 (3rd quarter)	-0.09 (-3.27)*	-0.09 (-3.31)*	-0.04 (-1.12)	-0.05 (-1.49)
SEA_3 (4th quarter)	-0.10 (-4.59)*	-0.10 (-4.60)*	-0.03 (-1.47)	-0.04 (-1.73)
Intercept	0.11 (5.16)*	0.11 (5.36)*	0.03 (1.41)	0.04 (1.70)
R^2	0.82	0.82	0.84	0.81

Note:

1. Numbers in parentheses are *t*-values, and * indicates significance at the 5% levels.
2. The hypothesis of no serial correlation cannot be rejected at the 5% level; Variables are in logarithmic terms with detrending.
3. To mitigate collinearity between U.S. and Canadian capacity in both datasets (0.94), the equations were estimated with either U.S. or Canadian capacity and reported the estimation that gives a better fit. Similarly, disposal income and wage rates were dropped.

4.3.2.1. Results of Lumber Price Model

With dataset I, the final model for lumber prices is a function of housing starts, mortgage rates, U.S. and Canadian wage rates, U.S. production capacity, Canadian log prices, exchange rates, seasonality, and dummies for the trade restrictions and log market shocks. Note that to mitigate collinearity between U.S. and Canadian capacity (0.94) in both datasets, the RFE models were estimated with either U.S. or Canadian capacity, and the estimation that gives a better fit was selected. The same procedure was applied to between disposable income and mortgage rates (correlation:-0.93). Housing starts, U.S. wage rates and production capacity, Canadian log prices, and exchange rates are statistically significant at the 5% level. For dummy variables, the MOU is statistically significant, pushing up the lumber price by 13% for the period 1986:4-1991:3. However, a single and multiple dummies for the SLA had little impact on lumber prices. Finally, DUM_1 for log market shock is statistically significant, suggesting that the federal harvest reductions pushed up lumber prices by 15%.

With dataset II, the final model for lumber prices was a function of housing starts, mortgage rates, U.S. wage rates, log prices, U.S. production capacity, exchange rates, seasonality, and dummies for the trade restrictions. Note that to mitigate collinearity between U.S. and Canadian wage rates (-0.74) in both datasets, the RFE models were estimated with either U.S. or Canadian wage rates, and the estimation that gives a better fit was selected. The same procedure was applied to between U.S. and Canadian log prices (correlation:-0.83). Housing starts and log prices are statistically significant at the 5% level. In addition, a single and multiple dummies for the SLA are not statistically significant. It is thus proven that the SLA had little impact on lumber prices.

4.3.2.2. Results of Canadian Export Supply Model

With dataset I, the final model for Canadian export supply was a function of housing starts, mortgage rates, U.S. and Canadian wage rates, Canadian production capacity, Canadian log prices, exchange rates, seasonality, and dummies for the trade restrictions and log market shocks. Housing starts, U.S. wage rates and Canadian production capacity, Canadian wage rates, and exchange rates are statistically significant at the 5% level. For dummy variables, the MOU is statistically significant, and it pushed down Canadian exports by 13% for the period 1986:4-1991:3. However, as in the RFE model for lumber prices, a single and multiple dummies for the SLA had little impact on Canadian exports. In addition, unlike the RFE model for lumber prices, dummies for log market shock are not statistically significant.

With dataset II, the final model for lumber prices was a function of housing starts, mortgage rates, U.S. wage rates, log prices, U.S. production capacity, exchange rates, seasonality, and dummies for the trade restrictions. Like the RFE model for lumber prices, only housing starts and log prices are statistically significant at the 5% level. However, a single and multiple dummies for the SLA are not statistically significant. Interestingly, the MOU is significant, and it depressed Canadian exports by 6 %.

4.3.2.3. Robustness of RFE Model in Previous Studies

The RFE model estimated by OLS is the best linear unbiased estimator if the Gauss-Markov (G-M) assumptions are all met. However, if any one of the assumptions fails, then OLS estimators are not unbiased (Wooldridge 2000). As such, robust estimation involves construction of alternative estimators, which are less sensitive to violations of the assumption than is the OLS estimator (Darnell 1994). Although robust estimator could, in principle, be constructed for any violation of the G-M assumptions, the focus of robust estimation generally involves violations of the assumed condition for zero mean and zero correlation. Zero mean condition involves the error (u) has an expected value of zero, or $E(u) = 0$, given any values of the independent variables. Zero correlation condition means that correlation between the error term (u) and independent variables are zero, or $Cov(x_i, u) = 0$, for $i = 1,2,..,k$, where x_i are independent variables. Thus, correlation between u and any of independent variables generally causes all of the OLS estimators to be inconsistent (Wooldridge 2000).

The inconsistency in OLS estimators is closely related to the problem of excluding a relevant variable (omitted variable bias) caused by the attribute of an RFE. That is, since an RFE is expressed as an endogenous variable in terms of exogenous variables in a single equation, multicollinearity problem usually arises. For example, in his RFE model, Zhang (2001) considered many exogenous variables as determinants of lumber prices, and then dropped several variables from the final model to avoid multicollinearity. In fact, when we meet multicollinearity problem in empirical analysis, it is reasonable to drop some independent variables from the model in an effort to reduce multicollinearity. In addition, high (but not perfect) correlation between independent variables is not a violation of the G-M Theorem. However, dropping a variable that belongs in the population model may cause the problem of omitted variable bias. As such, the OLS estimators can be inconsistent (Wooldridge 2000).

For example, this study can prove that excluding log prices in the RFE model could cause omitted variable bias, resulting in the inconsistency in the OLS estimators.

Specifically, first, this study regressed log prices on other independent variables as follows.

$$LP = f(HS, DI, MR, PC, W, CLP, CPC, CW, EX, MOU, SLA)$$

Then, let γ denote residual from this regression. Second, using the property of probability limits, correlation between γ and u — residual from the RFE model for lumber prices — can be obtained.[18] As a result of the estimation, if γ and u are not correlated, then the OLS estimators without including log prices in the model are consistent. If γ and u are positively correlated, the OLS estimators without including log prices are then overestimated. Or, if γ and u are negatively correlated, the OLS estimators without including log prices are then underestimated.

From these two steps, this study found that γ and u are positively correlated — around 0.3 and 0.2 for lumber price and Canadian export models, respectively. It implies that excluding log prices in the RFE model cause the omitted variable bias, which in turn tend to overestimate the effects of the trade restrictions. As a result, it is essential to incorporate log prices in determining lumber price/volume changes due to the trade restrictions, or at least introduced dummy variables to capture effects of the market shocks in the early 1990s. However, previous studies that estimated the impacts of the trade restriction, particularly the SLA (Zhang 2001), have never included log prices in their RFE models, which must have caused their OLS estimators to be inconsistent and biased.

[18] The property of probability limits can be defined as follows.

$P\lim \tilde{\beta}_i = \beta_i + Cov(x_i, u) / Var(x_i)$, where $\tilde{\beta}_i$ is estimated coefficients of independent variables (x_i). If the estimators are consistent, $Cov(x_i, u) = 0$ (Wooldridge 2000).

4.4. Summary and Conclusions

With two datasets, this study used the procedure of the 3SLS to estimate the SSE model in linear functional forms. The results show that the price effects of U.S. demand for and supply of lumber, and Canadian export supply are generally smaller than those commonly used by previous studies. Specifically, the price elasticities are -0.08~-0.13 for domestic demand, 0.15~0.26 for domestic supply, and 0.13~0.15 for Canadian exports, respectively. In contrast, other market factors, such as housing starts, log prices, and exchange rates, have significant impacts on U.S. demand and supply for lumber, and Canadian export supply. For instance, the price elasticity of demand is the smallest among all of the factors affecting lumber demand: housing starts, lumber prices, mortgage rates, and disposable income. Similarly, the price elasticity of domestic supply is smaller than that of wage rates and production capacity, and the price elasticity of Canadian exports is smaller than that of exchange rates and Canadian production capacity. As a result, other market factors such as housing starts, log prices, and exchange rates, have played key roles in shaping the U.S. lumber market. Particularly, the dynamics of log market have affected domestic and export supply significantly. For example, the federal harvest reductions in the PNW caused the domestic production to decrease by 5.4%, resulting in lumber prices and volume changes. Finally, the SSE model found that while the MOU depressed Canadian exports by 8%, the SLA had little impact on Canadian exports.

Now, this study can compare the price elasticities obtained from the SSE model with those of previous studies. Price elasticities used in previous studies are summarized in Table 4.5. In Wear and Lee (1993) and Zhang (2001), the price elasticity of demand for softwood lumber (–0.174) came from Adams et al. (1986), Spelter (1992), and Adams and Haynes (1996). Similarly, the elasticity of supply (0.4) was used as an average of the three regional counterparts for the Pacific, Interior, and South (0.239, 0.460, and 0.510), as documented in Adams et al. (1986). For the price elasticity of Canadian export supply, Zhange's (2001) 0.9 was an average of the estimates reported by Adams et al. (1986, 0.917), and Boyd and Krutilla (1987, 0.89), while Wear and Lee (1993) assumed unity for analytic convenience. The price elasticities obtained in this study are similar to those of Myneni et al. (1994). They estimated price elasticities derived from their SSE model for demand, supply and Canadian export supply for the U.S. market are –0.10, 0.27, and 0.19, respectively.

Based on the RFE models for lumber prices and Canadian export supply, the MOU might have pushed up the lumber price by 11-13%, and caused the Canadian export supply to drop by as much as 13%. These significant REF effects of the MOU are far greater than that obtained from the SSE model. As to the SLA, unlike the result of Zhang (2001), however, it had little impact on lumber prices or Canadian export supply. It implies that Zhang must have overestimated his SLA effect on lumber prices, due mainly to the omitted variable bias.

Table 4.5. Sources of price elasticities for softwood lumber used in previous studies

Source	DEM	SUP	EXP
Adams et al. (1986)		0.239 (P), 0.460 (I), 0.510 (S)	0.917
Boyd and Krutilla (1987)			0.890
Spelter (1992)	-0.174		
Myneni et al. (1994)	-0.10	0.27	0.19
Adams and Haynes (1996)	-0.174	0.572 (P), 0.574 (I), 0.950 (S)	0.625

Note:
1. P, I, and S represent Pacific, Interior and South, respectively.

CHAPTER 5

A NONSTRUCTURAL MODEL FOR U.S. SOFTWOOD LUMBER MARKET

This chapter provides empirical analyses for the nonstructural models, including the Johansen co-integration test and the vector error-correction (VEC) model, for the U.S. softwood lumber market. The main tasks of the nonstructural modeling are not only to discern the dynamic linkages among related market variables in the short run and long run, but also to assess the impacts of the U.S. restrictions in a corroborative way. For this purpose, the Johansen multivariate co-integration test was first introduced to estimate a long run nonstructural model on time-series. Long run is defined as a state of equilibrium in which economic forces such as demand and supply are in balance and thus there is no tendency to change. When it comes to dealing with non-stationary data series in an econometric model, therefore, the long-run equilibrium is identical to the concept of co-integration (Harris 1995). Non-stationary time-series data can be said co-integrated, if a linear combination of them is stationary, which thus tends to move closely together over time (Wooldridge 2000). Second, the VEC model was employed to estimate the short run dynamics in the relationship among time-series variables. Short run is defined as a state of disequilibrium in which adjustment to the equilibrium is occurring. The short run model is closely associated with the VEC model, which provides the useful information on the short-run adjustment behavior of economic variables (Wooldridge 2000). The two models are closely linked since the co-integration relationships are explicitly included in the VEC model. This chapter is organized as follows. First, the theoretical framework for the Johansen co-integration test and the VEC model is presented. Next, the empirical results as well as the estimation procedures for the two models are discussed. Finally, this chapter ends with a summary and conclusions.

5.1. Theoretical Framework

5.1.1. Johansen Co-integration Test

If a series is differenced d times before it becomes stationary, then it is integrated of order d, denoted as $I(d)$ (Harris 1995). Consider two time-series, y_t and x_t, which are both $I(d)$. In general, any linear combination of the two series will be $I(d)$. In other words, the residual obtained from regressing y_t on x_t, say $u_t = y_t - \beta x_t$, will be $I(d)$. However, if it is possible that for some $\beta \neq 0$, $u_t = y_t - \beta x_t$ is of a lower order of integration, $I(d-b)$, where $b > 0$, the two time-series, y_t and x_t, can be said to be co-integrated of order $(d-b)$ (Engle and Granger 1987). If y_t and x_t are co-integrated, a regression of y_t on x_t indicates that there is a long-run equilibrium relationship between y_t and x_t (Wooldridge 2000). In other words, the two time-series will move closely together over time and the difference between them will be stable, or stationary. Thus, the concept of co-integration indicates the existence of a long-run equilibrium to which an economic system converges over time (Harris 1995).

The Johansen co-integration test procedure is a multivariate method with simultaneous maximum likelihood estimation (Johansen and Juselius 1990, Johansen 1995). Given a vector y_t of n potentially endogenous variables, it is possible to model y_t as an unrestricted vector auto-regression (VAR) involving up to k lags:

$$y_t = A_1 y_{t-1} + ... + A_k y_{t-k} + u_t \quad (5.1)$$

where A_k is an $(n \times n)$ matrix of parameters, and u_t is a vector of normally and independently distributed error terms, or white noise. Sims (1980) has advocated this type of VAR-model as a way to estimate dynamic relationships among jointly endogenous variables without imposing *a priori* restrictions. In other words, unlike the structural econometric model discussed in the previous chapter, a VAR system does not require particular structural relationships and/or the exogeneity of some variables in the system. The system is in reduced form with each variable y_t regressed on only lagged variables of both itself and all the other variables in the system. Therefore, ordinary least squares (OLS) is efficient to estimate each equation because the right-hand side of each equation in the system consists of a common set of regressors including lagged and predetermined variables (Harris 1995 and Wooldridge 2000).

The equation (5.1) can be reformulated as follows:

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-k} + u_t \quad (5.2)$$

where $\Gamma_i = -(I - A_1 - ... - A_i), (\mathrm{i} = 1,...,\mathrm{k}-1)$ and $\Pi = -(I - A_1 - ... - A_k)$. The system specified this way contains information on both short run and long run adjustments to changes in y_t, via the estimates of Γ_i and Π, respectively. That is, $\Pi = \alpha\beta'$, where α represents the speed of adjustment to equilibrium and β' is a matrix of long-run coefficients such that the term $\beta' y_{t-k}$ represents up to $(n-1)$ co-integration relationships in the system (Johansen 1995).

Assume that y_t is a vector of non-stationary $I(1)$ variables. Then, while all the terms involving Δy_{t-k} in (5.2) are $I(0)$, Πy_{t-k} also should be stationary, or $I(0)$, for

$u_t \sim I(0)$ to be white noise (Hamilton 1995). There are three cases when the requirement of $\Pi y_{t-k} \sim I(0)$ is satisfied. First, when all the variables in y_t are stationary (Π has a full rank), then the variables in y_t are $I(0)$. Second, when there is no co-integration at all (Π has a zero rank), implying that there are no linear combinations of the y_t, which are $I(0)$, and consequently Π is a $(n \times n)$ matrix of zeros. However, neither of these two cases is particularly interesting. Third, when there exists up to $(n-1)$ co-integration relationships (Π has reduced rank), Πy_{t-k} is $I(0)$. In this case, $r \leq (n-1)$ co-integration vectors exist in β. That is, r columns of β consist of r linearly independent combinations of the variables in y_t along with $(n-r)$ non-stationary vectors. As a result, testing for co-integration is identical to a consideration of the rank of Π. In other words, co-integration tests are to identify the number of r linearly independent columns in Π (Harris 1995, p.79).

5.1.2. Vector Error-Correction (VEC) Model

Even though it is possible to estimate a long run nonstructural model on time-series using the Johansen co-integration test, it is also of interest to consider the short run dynamics in the relationship among the variables under consideration. Particularly, because a state of equilibrium may rarely be observed, the short run structure of the model provides the useful information that can be obtained from the short-run adjustment behavior of economic variables (Wooldridge 2000). A state of disequilibrium is mainly attributed to the inability of economic agents to adjust to new information instantaneously, due mainly to substantial adjustment costs. As a result, in an econometric model, adjustment costs cause the current value of the dependent variable, say Y, to be determined not only by the current value of explanatory variables, say X, but also by past values of explanatory variables. In addition, when Y changes over time in response to current and previous values of X, lagged values of Y also enter the short run dynamic model. This inclusion of lagged values of Y on the right-hand of the equation is a means of simplifying the form of the dynamic model (Harris 1995).

If all variables in a vector stochastic process y_t are $I(1)$ and co-integrated, it is possible to formulate a short run dynamic model, which is known as a vector error-correction (VEC) model (Engle and Granger 1987).

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \ldots + \Gamma_{k-1} \Delta y_{t-k+1} + \alpha(\beta' y_{t-1}) + u_t \tag{5.3}$$

where $\beta' y_{t-1}$ is a measure of the error or deviation from the equilibrium, which is stationary since the series are co-integrated.

A comparison between equations (5.2) and (5.3) shows that the VEC model is closely linked to the concept of co-integration. Indeed, Engle and Granger (1987) show that if y_t and x_t are co-integrated, then there must exist a VEC model. Conversely, they show that a VEC model generates co-integrated series. In addition, because variables are co-integrated, the VEC model incorporates both short-run and long-run effects, which is known as a restricted VAR model (Doornik and Hendry 1994). That is, if at any time the equilibrium holds, $\beta' y_{t-1} = 0$. During periods of disequilibrium, this term is non-zero and measures the distance the system is away from equilibrium during time t. Hence, an estimate of α provides information on the speed of adjustment, which implies how the variable y_t changes in response to disequilibrium.

5.2. Johansen Co-integration Test

The main task of the Johansen co-integration test is to find a correctly specified VAR model to identify structural economic relationships underlying the long run model. For this purpose, with all non-stationary variables,[19] the Johansen co-integration test was implemented twice. The first co-integration test was applied to find how the all the non-stationary series are correlated and which of them can be treated as weakly exogenous (Johansen 1995). This is called an unrestricted VAR model. Then, another Johansen co-integration test was employed to the endogenous variables to identify the long run relationships among them.

Before implementing the Johansen co-integration test, the appropriate lag-lengths were first determined to define a correctly specified VAR model. For this purpose, a number of VAR lag selection criteria and diagnostic tests were adopted. The VAR lag selection criteria include Log Likelihood Ratio (LR), Final Prediction Error (FPE), Akaike Information Criterion (AIC), Schwarz Information Criterion (SC), and Hannan-Quinn Information Criterion (HQ, Quantitative Micro Software 1998). In addition, the diagnostic tests include the tests of the normality, serial correlation, and homoskedasticity for the estimated residuals. Note that even though the main task of the procedure is to find a correctly specified VAR model, this is not always possible. For this reason, the minimum requirement set for an appropriate VAR model is that the model selected by the lag selection criteria is free of serial correlation in diagnostic tests (Doornik and Hendry 1994).

[19] Stationary series such as U.S. and Canadian lumber capacity were excluded from the analysis since they are not suited to the Johansen co-integration test.

5.2.1. Dataset I

The order of the VAR specification was first chosen on the basis of the five lag selection criteria. VAR of order 1 through 5 were investigated in logarithmic forms.[20] As shown in Table 5.1, the results are mixed. For example, while the FPE, AIC and HQ tests indicate 5 lags, LR and SC tests indicate 4 lags and 1 lag, respectively.

Table 5.1. Results of lag order selection criteria (dataset I)

Lag	LR	FPE	AIC	SC	HQ
0	NA	3.30E-20	-16.48	-16.19	-16.36
1	1416.08	1.07E-27	-33.74	-30.53*	-32.45
2	202.12	5.05E-28	-34.59	-28.47	-32.13
3	162.90	3.20E-28	-35.31	-26.28	-31.68
4	183.28*	7.99E-29	-37.26	-25.32	-32.46
5	124.27	5.42E-29*	-38.74*	-23.88	-32.77*

Note:
1. * indicates lag order selected by criterion at 5% level.
2. LR, FPE, AIC, SC, and HQ represent log likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion, respectively.

The diagnostic tests were then examined with 1, 4, and 5 lags to find the most appropriate approximation of the data generating process (DGP). The 5th order VAR specification was chosen because the VAR satisfied the minimum requirement for the correctly specified model, which is free of serial correlation. Table 5.2 summarizes the results of diagnostic tests. In the residual serial correlation test for individual equations and in the whole system, the null hypothesis of no serial correlation was accepted at the 5% significant level. In addition, the null hypothesis of no heteroskedasticity was accepted at the 5% significant level for all the equations. However, normality is rejected in two equations (EXP and CLP) and in the system.

[20] With the sample size ($n = 88$), the software (EViews) allowed up to 5 lags for the tests.

Table 5.2. Results of residual tests (dataset I)

Equation	Serial Correlation F_{AR} (2,27)	Heteroskedasticity F_{ARCH} (4,21)	Normality χ^2(2)
ΔEXP	0.65 [0.53]	0.10 [0.98]	26.24 [0.00]*
ΔSUP	0.76 [0.47]	1.42 [0.25]	0.95 [0.62]
ΔP	0.95 [0.39]	0.63 [0.65]	2.53 [0.28]
ΔHS	1.12 [0.34]	0.98 [0.43]	1.01 [0.60]
ΔDI	1.98 [0.15]	0.51 [0.73]	1.69 [0.43]
ΔMR	0.01 [0.99]	1.47 [0.23]	1.01 [0.60]
ΔW	2.02 [0.14]	1.19 [0.33]	0.95 [0.62]
ΔCW	0.65 [0.53]	0.46 [0.77]	0.41 [0.82]
ΔEX	0.52 [0.10]	0.47 [0.76]	1.91 [0.39]
ΔCLP	0.32 [0.73]	3.58 [0.17]	26.69 [0.00]*
System	0.04 [0.99]	-	65.67 [0.00]*

Note:

1. Values in brackets are p-value; * indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LM test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

The 5th order VAR model was tested by using the Johansen co-integration procedure. If the series were co-integrated, then it would be possible to obtain some indication of the weak exogeneity of the variables (Johansen 1995). Note that because one problem with the VAR model is its high dimension, or many variables, weak exogeneity with respect to the long-run parameters allows the co-integrated relations to be estimated from systems of smaller dimension than the whole model (Hetemaki et al. 2001). The results of the Johansen co-integration test are reported in Table 5.3, which indicates that there are two co-integration equations at the 5% significant level, or rank two $(r = 2)$. With this knowledge, this study tested the weak exogeneity and found seven weakly exogenous variables, which are HS, DI, MR, W, CW, EX, and CLP (Table 5.4). As a result, the VAR model for dataset I is defined as three endogenous variables (EXP, SUP, and P) and seven exogenous variables.

Table 5.3. Results of Johansen co-integration test (dataset I)

Null hypothesis	Eigenvalue	Trace Statistic	5% Critical Value
R=0	0.692	360.05	233.13*
R≤1	0.615	261.09	192.89*
R≤2	0.489	150.15	156.00
R≤3	0.423	121.10	124.24
R≤4	0.242	74.90	94.15
R≤5	0.223	51.65	68.52
R≤6	0.177	30.45	47.21
R≤7	0.086	14.05	29.68
R≤8	0.070	6.46	15.41
R≤9	0.004	0.37	3.76

Note

1. * denotes rejection of the null hypothesis at the 5% level.
2. With 5 lags, the test indicates 2 co-integration equations at 5% level ($r = 2$).

Table 5.4. Results of weak exogeneity tests (dataset I)

Equations	Weak exogeneity $H_0 : \alpha_i = 0$
EXP	21.07 [0.00]*
SUP	11.33 [0.00]*
P	7.53 [0.00]*
HS	3.03 [0.08]
DI	2.64 [0.11]
MR	0.29 [0.59]
W	0.99 [0.32]
CW	0.74 [0.39]
EX	0.83 [0.37]
CLP	0.63 [0.42]

Note:

1. * denotes rejection of the hypothesis at the 5% level.
2. Weak exogeneity under $r = 2$ follows χ^2 (2).

Another Johansen procedure was applied to identify the long-run co-integration relationships among the endogenous variables. That is, with the three endogenous variables (EXP, SUP, and P), five lag selection criteria and diagnostic tests were adopted to construct the correctly specified VAR model. VAR of order 1 through 5 were examined in logarithmic forms. As shown in Table 5.5, the results are mixed. For example, LR and the FPE tests indicate 4 lags. While AIC tests indicate 5 lags, SC and

HQ tests show 1 lag. The diagnostic tests were then examined with 1, 4, and 5 lags to find the most appropriate approximation of the data generating process (DGP). As a result, the 4th order VAR specification was chosen due to no serial correlation (Table 5.6).

With the 4th order restricted VAR model, the Johansen co-integration procedure was implemented to identify possible long run relationships among the three endogenous variables (EXP, SUP, and P). As shown in Table 5.7, the Johansen co-integration test indicates that there is only one co-integration equation at 5% significant level, or rank one ($r = 1$).[21] Figure 5.1 shows the time-series of co-integration vectors and recursive estimates of the three non-zero eigenvalues. The first three graphs show the normalized β-matrix to see if any of the $\beta' y_t$ are stationary. The next three graphs show long run fitted and actual values. The first co-integration vector looks fairly stationary, with fitted and actual values tracking each other reasonably closely. The last two vectors seem less non-stationary (upward trending), with not much relation between fitted and actual values. In addition, the last three graphs indicate that the eigenvalues are relatively constant, the first one at non-zero value, and the last two at close to zero throughout. As a result, these plots satisfy there is one co-integrated relationship among the three endogenous variables.

[21] The Johansen co-integration test included seven exogenous variables and seasonal dummy variables. However, the system specification tests indicated that trend was unnecessary.

Table 5.5. Results of lag order selection criteria for endogenous variables (dataset I)

Lag	LR	FPE	AIC	SC	HQ
0	NA	1.97E-05	-2.32	-1.97	-2.18
1	349.18	2.47E-07	-6.70	-6.09*	-6.46*
2	21.58	2.29E-07	-6.78	-5.91	-6.43
3	8.14	2.54E-07	-6.68	-5.54	-6.22
4	22.71*	2.26E-07*	-6.80	-5.40	-6.24
5	14.72	2.24E-07	-6.82*	-5.15	-6.15

Note:

1. * indicates lag order selected by criterion at 5% level.
2. LR, FPE, AIC, SC, and HQ represent log likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion, respectively.

Table 5.6. Results of residual tests for endogenous variables (dataset I)

Equation	Serial Correlation $F_{AR}(5,59)$	Heteroskedasticity $F_{ARCH}(4,56)$	Normality $\chi^2(2)$
ΔEXP	2.24 [0.07]	0.40 [0.81]	29.83 [0.00]*
ΔSUP	1.01 [0.42]	2.14 [0.09]	7.15 [0.03]*
ΔP	1.48 [0.21]	0.25 [0.91]	5.38 [0.07]
System	1.19 [0.22]	-	54.85 [0.00]*

Note:

1. Values in brackets are p-value; * indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LS test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

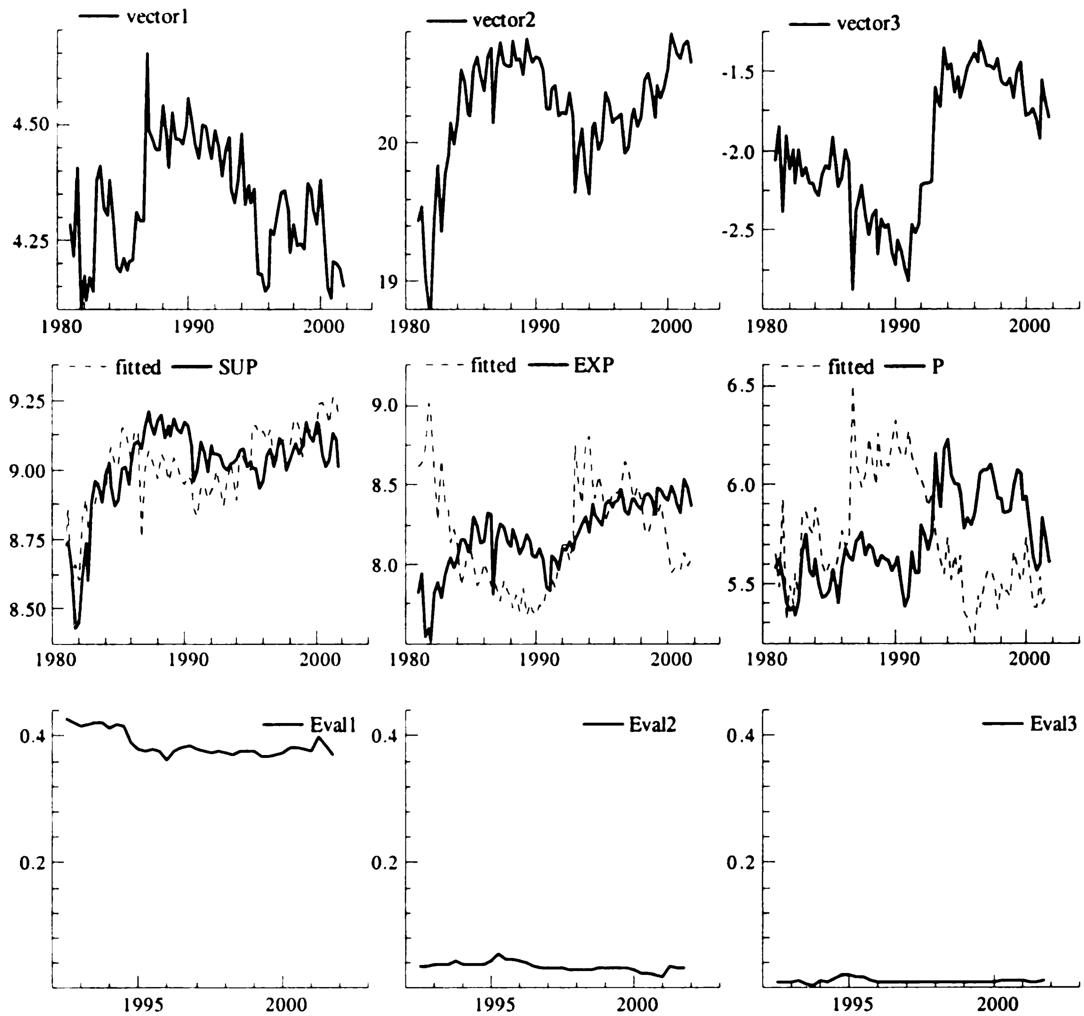
Table 5.7. Results of Johansen co-integration test for endogenous variables (dataset I)

Null hypothesis	Eigenvalue	Trace Statistic	5% Critical Value
R=0	0.370	42.34	29.68*
R≤1	0.031	3.54	15.41
R≤2	0.011	0.91	3.76

Note:

1. * denotes rejection of the null hypothesis at the 5% level.
2. With 4 lags, the test indicates 1 co-integration equations at 5% level ($r = 1$).

Figure 5.1. Co-integration vectors and recursive estimates of the eigenvalues (dataset I)



5.2.2. Dataset II[22]

VAR of order 1 through 4 were considered in logarithmic forms (Table A-1). The results of dataset II are also mixed. That is, the LR, FPE, and AIC tests indicate 3 lags, whereas SC and HQ tests indicate 4 lags. The diagnostic tests were then examined with 3 and 4 lags to find the most appropriate VAR model and thus the 3rd order VAR specification was chosen due to no serial correlation. All the diagnostic tests for individual equations and in the whole system were accepted at the 5% significant level (Table A-2).

The Johansen co-integration test was implemented to examine the 3rd order VAR model (Table 5.8). The results indicate that there are two co-integration equations at 5% significant level, or rank two $(r = 2)$. In addition, the weak exogeneity test in the co-integration vector found six weakly exogenous variables, which are HS, DI, MR, W, EX, and LP (Table A-3).

Table 5.8. Results of Johansen co-integration test (dataset II)

Null hypothesis	Eigenvalue	Trace Statistic	5% Critical Value
R=0	0.757	240.61	192.89*
R≤1	0.596	168.27	156.00*
R≤2	0.521	122.09	124.24
R≤3	0.455	84.51	94.15
R≤4	0.352	53.56	68.52
R≤5	0.300	31.47	47.21
R≤6	0.161	13.25	29.68
R≤7	0.072	4.32	15.41
R≤8	0.010	0.51	3.76

Note:
1. * denotes rejection of the null hypothesis at the 5% level;
2. With 3 lags, the test indicates 2 co-integration equations at 5% level ($r = 2$).

[22] Due to the small sample size ($n = 52$), the nonstructural modeling with dataset II confined to the U.S. market variables. Hence, among non-stationary variables, Canadian wage rates (CW) and Canadian log prices (CLP) were excluded from both the co-integration test and VEC modeling.

Another Johansen method was applied to identify the long-run co-integration relationships for the three endogenous variables. VAR of order 1 through 4 were examined in logarithmic forms (Table A-4). The results show that while the LR, AIC and SC tests indicate 2 lags, the FPE and HQ tests indicate 3 lags. The diagnostic tests were then examined with 2 and 3 lags to find the most appropriate VAR model, and thus the 2[nd] order VAR specification was chosen due to no serial correlation. All the diagnostic tests for individual equations and in the whole system were accepted at the 5% significant level (Table A-5).

With the 2[nd] order restricted VAR model, the Johansen co-integration procedure was implemented to identify possible long run relationships among the three endogenous variables (EXP, SUP, and P). The Johansen co-integration test (Table 5.9) indicates that there is only one co-integration equation at the 5% significant level, or rank one ($r = 1$)[23].

Table 5.9. Results of Johansen co-integration test for endogenous variables (dataset II)

Null hypothesis	Eigenvalue	Trace Statistic	5% Critical Value
R=0	0.366	33.66	29.68*
R≤1	0.174	11.81	15.41
R≤2	0.053	2.61	3.76

Note:
1. * denotes rejection of the null hypothesis at the 5% level.
2. With 2 lags, the test indicates 1 co-integration equations at 5% level ($r = 1$).

As seen in dataset I, plots of the co-integration vectors and recursive estimates of eigenvalues are presented in Table B-1. These indicate that while the first two vectors are

[23] The Johansen co-integration test included six exogenous variables and seasonal dummy variables. However, the system specification tests indicated that trend was unnecessary.

non-stationary, the third looks stationary. In addition, the eigenvalues are relatively stable over time, the first one at non-zero value, and the last two at close to zero throughout. Thus, these plots satisfy there is one co-integrated relationship among the three endogenous variables.

5.2.3. Testing for Unique Co-integration Vectors

After having determined how many co-integration vectors are in the VAR models, it is now necessary to consider whether these are unique and thus whether they tell us anything about the structural economic relationships underlying the long run model (Harris 1995). In general, because the Johansen co-integration procedure only determines how many unique co-integration vectors span the co-integration space, the estimates produced for any particular column in the co-integration vectors, β, are not necessarily unique. As a result, it is necessary to impose linear restrictions on the co-integration vectors to obtain unique vectors lying within the co-integration space, and then test whether the co-integration vectors are identified. This is called testing for identification (Doornik and Hendry 2001).

To obtain unique co-integration vectors for the VAR models, the implementation of the restriction tests proceeded as follows. First, with three endogenous variables (EXP, SUP, and P) in the VAR models, two coefficients were restricted at a time to equal 1 and –1 and the third to equal an unrestricted value (denotes φ) in β vector such as $\beta = [1, -1, \varphi]$ to see if particular relationships span the co-integration space. Then, joint tests of restriction on both the co-integration vectors (β) and speed-of-adjustment parameters (α) were implemented to consider weakly exogenous to the long run parameters in the restrictive models. Notice that this study found that there is only one co-integration vector in the VAR models. In general, when rank is one ($r = 1$), the co-integration space is uniquely defined by a single vector. However, the testing for unique co-integration vectors still requires because it may not be possible to identify unique vectors, and this does not invalidate the long run stationary relationships between the variables in the co-integration space (Harris 1995).

For the VAR model with dataset I, only the restriction of the coefficients for Canadian export supply (EXP) and domestic supply (SUP) at a time to equal 1 and –1, and lumber price (P) to equal an unrestricted value ($\beta = 0.27$) was not rejected since the LR test statistic is not significant ($\chi^2(1) = 0.14$). Then, with the test of the restrictions being $\chi^2(2) = 5.36$, the linear restriction and the feedback to zero for P related to long run weak exogeneity were only accepted. Similarly, with dataset II, the restriction of the coefficient for EXP and SUP at a time to equal 1 and –1, and P to equal an unrestricted value ($\beta = 0.31$) was only accepted ($\chi^2(1) = 1.95$). Then, the linear restriction and the weak exogeneity test for P was only accepted ($\chi^2(2) = 3.60$). The results are reported in Table 5.10.

The results of testing for unique co-integration vectors in the VAR models indicate that there exist long run relationships between Canadian export supply and domestic supply. In addition, the joint tests show that U.S. lumber price is weakly exogenous to the long run parameters in the U.S. softwood lumber market. However, weak exogeneity for Canadian export supply and domestic supply was rejected. These findings will be used for the VEC modeling in the following section. Notice that the long run responses of Canadian export supply and/or domestic production to U.S. lumber prices are relatively small, which are 0.27-0.31. That is, when U.S. lumber prices increase by 1%, Canadian lumber exports to the U.S. only increases by 0.27-0.31% in the long run.

Table 5.10. Results of testing for unique co-integration vectors

Hypothesis	LR test statistic	
	Dataset I	Dataset II
$\beta_1 = 1$, $\beta_2 = -1$, $\beta_3 = \varphi$	0.14 [0.71]	1.95 [0.16]
$\beta_1 = 1$, $\beta_2 = \varphi$, $\beta_3 = -1$	12.27 [0.00]*	10.90 [0.00]*
$\beta_1 = \varphi$, $\beta_2 = 1$, $\beta_3 = -1$	34.40 [0.00]*	11.05 [0.00]*
$\beta_1 = 1$, $\beta_2 = -1$, $\beta_3 = \varphi$, $\alpha_1 = 1$	23.67 [0.00]*	13.18 [0.00]*
$\beta_1 = 1$, $\beta_2 = -1$, $\beta_3 = \varphi$, $\alpha_2 = 1$	19.26 [0.00]*	11.23 [0.00]*
$\beta_1 = 1$, $\beta_2 = -1$, $\beta_3 = \varphi$, $\alpha_4 = 1$	5.36 [0.07]	3.60 [0.17]

Note:

1. β_1, β_2, and β_3 represent the coefficients of the co-integration vectors for *EXP*, *SUP* and *P*, respectively. In addition, α_1, α_2, and α_3 represent the coefficients of the speed-of-adjustment parameters for *EXP*, *SUP* and *P*, respectively.
2. LR test statistic is based on the χ^2 distribution and parenthesis is p-value.
3. φ represents an unrestricted value and * denotes significance at the 5% level.

5.3. Vector Error-Correction Model

The main task of the VEC model is to find the information on the short-run adjustment behavior of time-series variables. For this purpose, this study explicitly included the co-integration relationships that were already obtained, and estimated the VEC models. The VEC modeling employed the Hendry approach of a general-to-specific procedure (Doornik and Hendry 1994, and Hendry 1995). First, this study constructed the basic VEC model by differencing and the co-integration relationships. Second, the parsimonious VEC (PVEC) model was estimated based on tests of the significance of the variables in the basic VEC model. In addition, the conditional PVEC (CPVEC) model was estimated by conditioning on any weakly exogenous variables in the system. Finally, the structural model was constructed to consider any simultaneous effects among the variables in the PVEC model, and test to ensure that the resulting restricted model parsimoniously encompasses the PVEC model.

5.3.1. Dataset I

5.3.1.1. Basic VEC Model

Having obtained the long run co-integration relations using the Johansen procedure, the basic VEC model was estimated with explicit inclusion of the error-correction terms. For this purpose, equation (5.3) is reformulated as:

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \ldots + \Gamma_{k-1} \Delta y_{t-k+1} + \alpha(\hat{\beta}' y_{t-1}) + u_t \tag{5.4}$$

where $\hat{\beta}' y_{t-1}$ is the error-correction term that consists of the co-integration relationship obtained from the Johansen approach. Based on equation (5.4), the basic VEC model was defined as follows:

$$\begin{bmatrix} SUP_t \\ EXP_t \\ P_t \end{bmatrix} = A_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \\ P_{t-1} \end{bmatrix} + \ldots + A_4 \begin{bmatrix} SUP_{t-4} \\ EXP_{t-4} \\ P_{t-4} \end{bmatrix} + B \begin{bmatrix} HS_t \\ DI_t \\ MR_t \\ W_t \\ CW_t \\ CLP_t \\ EX_t \end{bmatrix} + \Gamma \begin{bmatrix} MOU \\ SLA \\ DUM_1 \\ DUM_2 \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta}\, y_{t-1}) + \varepsilon_t$$

To be clear, this study has separated the normal exogenous variables from the dummies for trade restrictions, log market shocks, and seasonality. The lag length of endogenous variables ($k = 4$) is determined by the Johansen co-integration test. In addition, there is no restriction placed on the speed of adjustment (α) to encompass weakly exogenous. The results from estimating the basic VEC model are reported in Table 5.11. Notice that the error correction term for lumber prices was not significant in the basic VEC model.

To check whether the basic VEC model is consistent with the data evidence, the various diagnostic tests of the residuals for endogenous variables were implemented. The model fits the data quite well and has approximately white noise, normally distributed errors (Table B-2). Additionally, the first three graphs in Table B-3 show that the model has generally constant coefficients because the one-step residuals are placed within their approximate 95% confidence bands with constant standard errors. The next four graphs in Table B-3 then indicate that no break point Chow test is anywhere significant within the 1% line. As a result, no major specification problems were identified in the system (Table A-6).

Table 5.11. Results of the basic VEC model (dataset I)

Variable		EXP		SUP		P	
		Coefficient	*t*-value	Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.75	-4.86*	-0.29	-3.89*	-0.03	-0.29
	EXP_{t-2}	-0.57	-3.55*	-0.20	-2.98*	0.07	0.36
	EXP_{t-3}	-0.48	-3.21*	-0.20	-2.12*	-0.01	-0.97
	EXP_{t-4}	-0.32	-2.66*	-0.17	-0.84	0.09	0.59
	P_{t-1}	-0.11	-0.94	0.01	1.01	-0.15	-1.02
	P_{t-2}	-0.26	-1.13	-0.05	-0.59	-0.31	-2.06*
	P_{t-3}	0.01	0.02	0.23	2.78*	0.22	1.46
	P_{t-4}	-0.05	-1.65	0.03	0.30	-0.14	-0.93
	SUP_{t-1}	-0.32	-2.30*	-0.32	-2.03*	0.41	1.45
	SUP_{t-2}	0.17	0.89	-0.34	-2.39*	0.04	1.36
	SUP_{t-3}	0.04	0.19	-0.24	-2.65*	-0.05	-0.77
	SUP_{t-4}	-0.20	-1.08	-0.04	-0.32	-0.03	-0.14
Exogenous	HS	0.27	2.40*	0.25	3.06*	0.16	0.46
	MR	-0.01	-2.10*	-0.01	-0.96	-0.01	-0.74
	DI	0.21	1.40	-0.23	-1.32	-0.36	-0.94
	W	-1.47	-2.07*	0.16	0.31	-0.22	-0.26
	CW	0.10	0.36	-0.15	-0.42	-0.36	-0.88
	CLP	-0.09	-0.76	-0.19	-1.72	-0.04	-0.33
	EX	0.26	2.00*	-0.01	-1.06	0.24	1.56
	MOU (1986:4-1991:3)	-0.08	-2.51*	0.04	2.30*	0.04	0.91
	SLA (1996:2-2001:1)	-0.03	-0.63	0.04	1.42	0.02	0.56
	DUM_1 (1991:1-1993:4)	0.01	0.24	-0.06	-2.25*	0.11	1.99*
	DUM_2 (1997:3-1998:4)	0.01	0.17	0.01	0.31	-0.02	-0.48
	SEA_1 (2^{nd} quarter)	0.12	3.16*	0.11	4.20*	0.13	2.79*
	SEA_2 (3^{rd} quarter)	0.08	1.33	0.15	4.00*	0.06	0.88
	SEA_3 (4^{th} quarter)	0.03	0.57	0.06	1.98*	0.01	0.25
Error-correction term		-0.74	-4.17*	-0.49	-6.27*	-0.29	-1.85

Note:

1. * indicates significance at the 5% level.
2. Based on the co-integration test, four lags were identified.

5.2.2.2. Parsimonious VEC Model

To estimated the parsimonious VEC model, all the insignificant lagged and exogenous variables in the basic VEC model were eliminated using an F - test (Table A-

7).[24] The null hypothesis that the omitted explanatory variables are zero was not rejected by the overall F - test ($F(21,161) = 1.07$, p - value = 0.37). The results from estimating the PVEC model are reported in Table A-8. Notice that the error-correction term for lumber prices still were not significant at the 5% level.

From the Johansen co-integration tests and the significance of the error correction term in the PVEC model, it is now possible to restrict any weakly exogenous in the system, known as a conditional PVEC (CPVEC) model (Harris 1995). In other words, from the co-integration tests, lumber prices were identified as a weakly exogenous variable in the system. In addition, the error correction term in the PVEC model was not significant. As a result, lumber price can be conditioned in the system as follows:

$$\begin{bmatrix} SUP_t \\ EXP_t \end{bmatrix} = A_0 P_t + A_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \end{bmatrix} + A_2 \begin{bmatrix} SUP_{t-2} \\ EXP_{t-2} \end{bmatrix} + A_3 \left[EXP_{t-3} \right] + B \begin{bmatrix} HS_t \\ EX_t \\ MOU \\ SLA \\ DUM_1 \\ DUM_2 \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta}\, y_{t-1}) + \varepsilon_t$$

Notice that as a conditioned variable, lumber prices are located in the right-hand side without lagged variables. At this stage, α is restricted as $\alpha_3 = 0$, where α_3 is the speed of adjustment for lumber prices. The resulting reduction to the CPVEC model was supported by the overall F-tests. That is, the null that the omitted explanatory variables are zero was not rejected ($F\,(36,157) = 1.30$, p - value = 0.25). Finally, the results from estimating the CPVEC model are presented in Table 5.12. Notice that estimates of the

[24] Insignificant variables must be dropped from the system altogether since all equations have the same formulation by construction (Doornik and Hendry 2001).

coefficients for the error-correction terms verify that the speed of adjustment to long run changes in the Canadian exports and U.S. production is relatively slow but significant.

Table 5.12. Results of the conditional PVEC model (dataset I)

Variable		EXP		SUP	
		Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.60	-4.63*	-0.10	-2.27*
	EXP_{t-2}	-0.38	-3.01*	-0.08	-2.16*
	EXP_{t-3}	-0.28	-2.62*	0.06	0.96
	SUP_{t-1}	-0.17	-2.39*	-0.33	-2.97*
	SUP_{t-2}	-0.10	-0.97	-0.30	-2.63*
	P	0.27	1.54	0.20	2.91*
Exogenous	HS	0.14	2.14*	0.09	2.03*
	EX	0.17	2.08*	-0.14	-1.43
	MOU (1986:4-1991:3)	-0.09	-3.67*	0.03	2.27*
	SLA (1996:2-2001:1)	-0.04	-1.38	0.01	0.45
	DUM_1 (1991:1-1993:4)	0.01	1.02	-0.04	-2.43*
	DUM_2 (1997:3-1998:4)	0.03	0.88	0.01	0.36
	SEA_1 (2nd quarter)	0.03	0.96	0.09	4.14*
	SEA_2 (3rd quarter)	0.09	2.11*	0.11	3.90*
	SEA_3 (4th quarter)	0.06	1.70	0.05	2.18*
Error-correction term		-0.56	-2.55*	-0.23	-2.91*
Multivariate tests		$F_{ar}(12,122) = 1.24$; $F_{het}(36,157) = 1.17$; $\chi^2(4) = 4.90$			

Note:

1. * indicates significance at the 5% level.
2. Based on the co-integration test, two lags were identified.
3. Multivariate tests for serial correlation, heteroskedasticity, and normality were not rejected at the 5% significant level.

5.2.2.3. Structural VEC Model

The main task for estimating the structural VEC model is to reduce simultaneous effects between endogenous variables in the (conditional) PVEC model (Doornik and Hendry 2001). For example, this study found that in the conditional PVEC model with dataset I, the residual of Canadian exports is strongly correlated with that of U.S. production (-0.42). This implies that the two equations may be simultaneously determined in the system, and thus OLS estimation for the model is no longer consistent. As such, the structural VEC model was estimated by imposing some restrictions on the CPVEC model. That is, according to economic theory, this study assumes that Canadian exports contemporaneously depend on the level of U.S. production in the lumber market. Thus, the structural VEC model included the endogenous variable for U.S. production as additional explanatory variables in Canadian export supply equation. In addition, from the CPVEC model, this study deleted all the lagged variables and exogenous variables with a t-statistic less than one (Doornik and Hendry 2001). The structural VEC model is now redefined as:

$$SUP_t = \Phi_0 P_t + A_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \end{bmatrix} + A_2 \begin{bmatrix} SUP_{t-2} \\ EXP_{t-2} \end{bmatrix} + B \begin{bmatrix} HS_t \\ EX_t \\ MOU \\ DUM_1 \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta} y_{t-1}) + \varepsilon_t$$

$$EXP_t = \Pi_0 P_t + \Pi_0 SUP_t + \mathrm{P}_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \end{bmatrix} + \mathrm{P}_2 [EXP_{t-2}] + \mathrm{P}_3 [EXP_{t-3}] + \mathrm{T} \begin{bmatrix} HS_t \\ EX_t \\ MOU \\ DUM_1 \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta} y_{t-1}) + \varepsilon_t$$

Notice that the endogenous variable for U.S. production (SUP_t) is included in the Canadian exports equation. As seen in the conditional PVEC model, the speed of adjustment (α) for lumber prices was restricted. The structural VEC model was estimated by full-information maximum likelihood (FIML) and reported in Table 5.13.

The resulting structural model was supported by the system reduction test. That is, the null hypothesis that the deleted explanatory variables are zero was not rejected (χ^2 statistic is 2.83, p - value = 0.59). More importantly, the LR test of over-identifying restrictions was not rejected ($\chi^2(4) = 2.52$, p - value = 0.64).[25] Thus, the structural VEC model can be said to parsimoniously encompass the conditional PVEC model.

The impulse responses for the SVEC are plotted in Figure 5.2. Impulse responses trace out the effect of a one-time shock to one standard deviation of the residual on current and future values of the endogenous variables —*SUP* and *EXP* (Doornik and Hendry 2001). The first column shows the effects of an unexpected one-time shock in domestic supply on all two endogenous variables. The second column shows the effect of an unexpected increase in Canadian exports on the endogenous variables. These estimated impulse responses show patterns of persistent common variation between domestic supply and Canadian exports. For example, an unexpected shock in Canadian exports on domestic supply slowly fades away over 5 quarters.

Finally, diagnostic tests of whether the model is congruent with the data evidence are provided in Tables B-4 and B-5. The structural VEC model appears to fit the data quite well, there is no evidence of within-equation residual serial correlation and the

[25] The structural model requires to be identified. The order condition for identification involves the number of unrestricted coefficients. Specifically, it specifies that no more than $l = (n \times (k-1)) + r$ explanatory variables can enter any equation and no more than $n \times l$ unknowns can enter the model (n = number of variables, k = lag lengths, and r = number of co-integration). In other words, the order condition ensures a sufficient number of equations, which can be identified by just counting the number of unrestricted coefficients. On the other hand, the rank condition requires that equations are linearly independent. Thus, the actual location of the restrictions placed on an unrestricted structural VEC model is an important element in identifying unique short run equations. As a general guide, each equation in the model requires at least one unique predetermined variables (entering with a non-zero coefficient) to identify it (Harris 1995 and Doornik and Hendry 2001). In this study, EXP_{t-3} only enters Canadian export equation, whereas SUP_{t-2} only enters U.S. production equation.

densities and distributions are close to normality (Table B-4). The model has generally constant coefficients and no break points in the individual equations and in the system as a whole (Table B-5). Moreover, the residual correlation between Canadian exports and U.S. production obtained from the FIML model is -0.07, which indicates some success in constructing the structural model. Figures 5.3 shows one-step static (*ex-post*) and dynamic forecast for the last four years of the data. None of the forecasts lie outside their individual confidence bars, and therefore constancy of the model is readily accepted. This is confirmed by the final graph, which shows the Chow test of stability, with all χ^2-values lying below 1% significance line. Finally, estimates of the coefficient of the error correction terms confirm that the speed of adjustment to long run changes in the variables is slow but significant.

Table 5.13. Results of the structural PVEC model (dataset I)

Variable		EXP		SUP	
		Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.61	-5.36*	-0.16	-2.89*
	EXP_{t-2}	-0.33	-2.98*	-0.15	-2.19*
	EXP_{t-3}	-0.29	-3.15*		
	SUP	-0.06	-2.29*		
	SUP_{t-1}	-0.18	-2.98*	-0.26	-3.20*
	SUP_{t-2}			-0.24	-2.94*
	P	0.28	1.83	0.18	2.94*
Exogenous	HS	0.20	3.16*	0.09	2.22*
	EX	0.25	2.57*	-0.15	-1.57
	MOU (1986:4-1991:3)	-0.09	-3.83*	0.03	2.97*
	SLA (1996:2-2001:1)	-0.04	-1.39		
	DUM_1 (1991:1-1993:4)	0.03	1.03	-0.04	-2.53*
	SEA_1 (2nd quarter)	0.04	1.16	0.10	5.40*
	SEA_2 (3rd quarter)	0.09	2.37*	0.10	5.35*
	SEA_3 (4th quarter)	0.06	1.74	0.04	2.01*
Error-correction term		-0.50	-2.97*	-0.23	-3.92*
Multivariate tests		$F_{ar}(12,124) = 1.44$; $F_{het}(36,160) = 1.28$; $\chi^2(4) = 1.84$			
LR test of over-identification		$\chi^2(4) = 2.52$			

Note:

1. * indicates significance at the 5% level.
2. Multivariate tests for serial correlation, heteroskedasticity, and normality were not rejected at the 5% significant level.
3. LR test of over-identification restriction was not rejected at the 5% significant level.

Figure 5.2. Impulse responses for SPVEC model (dataset I)

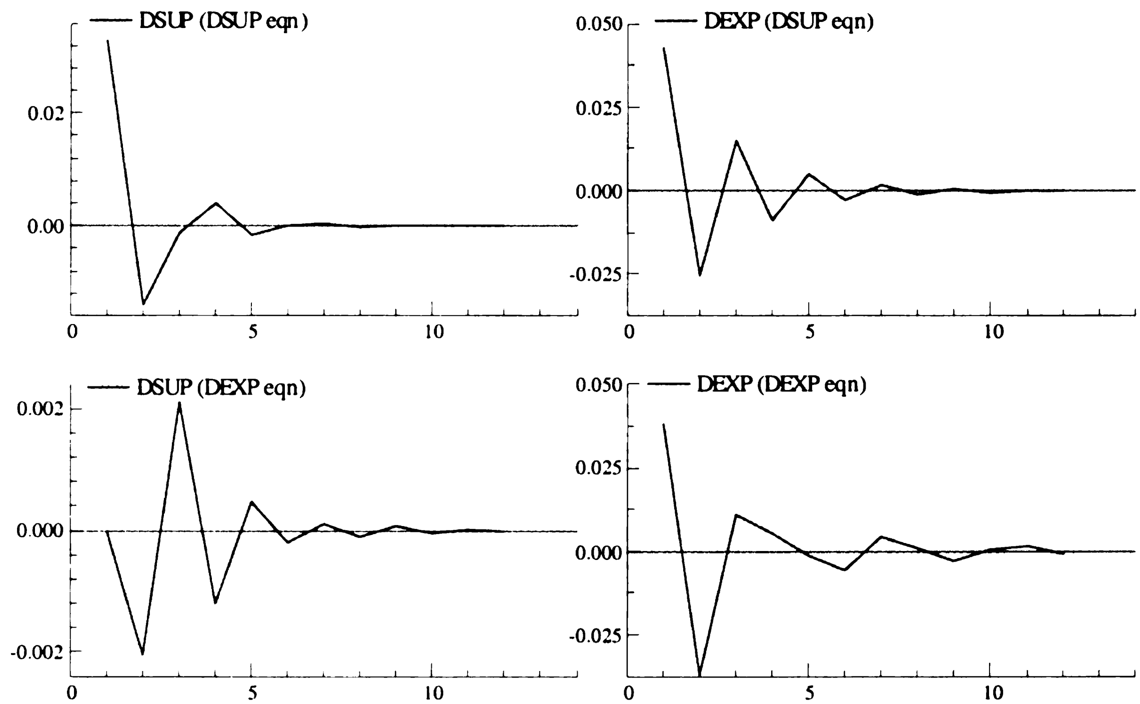
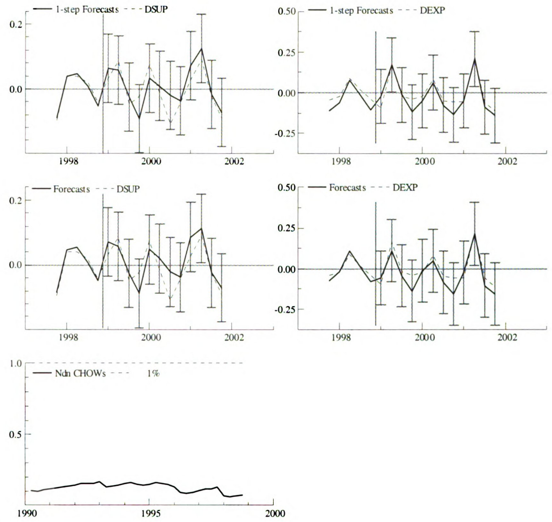


Figure 5.3. One-step and dynamic model-based forecasts for SPVEC model (dataset I)



5.3.2. Dataset II

5.3.2.1. Basic VEC Model

Based on equation (5.4), the basic VEC model for dataset II is defined as:

$$\begin{bmatrix} SUP_t \\ EXP_t \\ P_t \end{bmatrix} = A_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \\ P_{t-1} \end{bmatrix} + A_2 \begin{bmatrix} SUP_{t-2} \\ EXP_{t-2} \\ P_{t-2} \end{bmatrix} + B \begin{bmatrix} HS_t \\ DI_t \\ MR_t \\ W_t \\ LP_t \\ EX_t \end{bmatrix} + \Gamma \begin{bmatrix} MOU \\ SLA \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta}\, y_{t-1}) + \varepsilon_t$$

The lag length of endogenous variables ($k = 2$) is determined by the Johansen co-integration test. In addition, there is no restriction placed on the speed of adjustment (α) to consider weakly exogenous. Note that instead of the dummies for log market shocks, log prices is directly included in this system. The results from estimating the basic VEC model are reported in Table 5.14. Notice that unlike in dataset I, the error correction terms for the three endogenous variables were significant at the 5% significant level.

The diagnostic tests of the residuals for endogenous variables show that the model fits the data quite well and has approximately white noise, normally distributed errors (Table B-6). In addition, the model has generally constant parameters (Table B-7). As a result, no major specification problems were identified in the model (Table A-9).

Table 5.14. Results of the basic VEC model (dataset II)

Variable		EXP		SUP		P	
		Coefficient	*t*-value	Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.80	-5.15*	-0.50	-3.45*	-0.90	-3.85*
	EXP_{t-2}	-0.47	-2.51*	-0.43	-3.08*	-0.23	-1.24
	P_{t-1}	-0.07	-0.79	0.05	0.83	-0.07	-0.41
	P_{t-2}	-0.15	-1.46	0.04	1.20	-0.69	-3.41*
	SUP_{t-1}	-0.43	-3.90*	0.18	1.23	0.55	1.46*
	SUP_{t-2}	-0.18	-1.03	0.10	1.14	0.76	1.60
Exogenous	HS	0.28	2.09*	0.23	2.69*	1.19	1.27
	MR	-0.01	-0.13	0.02	1.68	-0.06	-1.38
	DI	0.26	0.64	-0.45	-1.50	-1.04	-1.30
	W	-0.19	-0.32	-0.40	-3.55*	0.87	1.28
	LP	0.14	2.46*	0.06	1.04	-0.12	-1.49
	EX	0.56	1.98*	-0.50	-0.70	-0.86	-2.91*
	MOU (1986:4-1991:3)	-0.03	-0.70	-0.04	-1.45	-0.09	-1.18
	SLA (1996:2-2001:1)	-0.04	-1.33	0.04	1.03	-0.04	-0.96
	SEA_1 (2^{nd} quarter)	0.09	2.02*	0.12	1.08	-0.04	-1.11
	SEA_2 (3^{rd} quarter)	0.14	3.03*	-0.05	-1.25	-0.09	-2.01*
	SEA_3 (4^{th} quarter)	0.09	2.40*	-0.01	-1.28	-0.09	-1.44
Error-correction term		-0.45	-4.44*	-0.42	-7.55*	-0.62	-2.52*

Note:
1. * indicates significance at the 5% level.
2. Based on the co-integration test, two lags were identified.

5.3.2.2. Parsimonious VEC Model

As in dataset I, this study deleted all the insignificant exogenous variables and lagged variables using an F-test (Table A-10). The null that the omitted variables are zero was not rejected by the overall F-test ($F(6,56) = 0.96, p\text{-value} = 0.46$). Thus, the PVEC model was reformulated based on the above procedures.

$$\begin{bmatrix} SUP_t \\ EXP_t \\ P_t \end{bmatrix} = A_1 \begin{bmatrix} SUP_{t-1} \\ EXP_{t-1} \end{bmatrix} + A_2 \begin{bmatrix} EXP_{t-2} \\ P_{t-2} \end{bmatrix} + B \begin{bmatrix} HS_t \\ LP_t \\ EX_t \end{bmatrix} + \Gamma \begin{bmatrix} MOU \\ SLA \\ Seasonality \end{bmatrix} + \alpha(\hat{\beta} y_{t-1}) + \varepsilon_t$$

The PVEC model were estimated and presented in Table A-11. Note that with dataset II, it was not appropriate to condition on lumber prices to construct the conditional PVEC model in view of the following facts. First, lumber prices in dataset II were also identified as a weakly exogenous from the co-integration involving the long run model. However, unlike dataset I, the error correction term in the equation was significant at the 5% level. Second, the resulting reduction to the conditional PVEC model was not supported by the overall F-test. That is, the null that the omitted explanatory variables are zero was rejected ($F(18,79) = 1.97$, p-value = 0.02). Thus, with dataset II, this study only estimated the PVEC model. Finally, the diagnostic tests indicate that there is no evidence against the specification (multivariate tests in Table A-11).

5.3.2.3. Structural VEC Model

Note that with dataset II, the residuals of the endogenous variables show little simultaneous effects. That is, the correlation coefficients between the three endogenous variables are less than 0.10. This study thus constructed the structural VEC model for dataset II by only imposing the restriction of weak exogeneity on the PVEC model. In addition, insignificant lagged variables were dropped from the PVEC model. Thus, the structural VEC model is estimated by full-information maximum likelihood (FIML) and reported in Table 5.15. The resulting structural VEC model was supported by the system reduction test. The null that deleted explanatory variables are zero was not rejected (χ^2 statistic is 1.26, p- value = 0.26). In addition, the LR test of over-identifying restrictions was not rejected ($\chi^2(6) = 6.33$, p - value = 0.27). As a result, as in dataset I, the structural VEC model can be said to parsimoniously encompass the PVEC model.

The impulse responses for the SVEC are plotted in Figure 5.4. The first column shows the effects of an unexpected one-time shock in domestic supply on all three endogenous variables. The second column shows the effect of an unexpected increase in Canadian exports on the endogenous variables. The last column shows the effect of a one-time shock in lumber prices on all three endogenous variables. As seen in dataset I, the estimated impulse responses between domestic supply and Canadian exports indicate patterns of persistent common variation. However, an unexpected rise in lumber prices on Canadian exports does not show patterns of persistent common variation in the short run (over two quarters).

The various diagnostic tests of the residuals for endogenous variables were implemented and provided in Tables B-8 and B-9. The tests show that the model fits the data well and has approximately white noise, normally distributed errors. The model has generally constant parameters. Additionally, Figure 5.5 show one-step static forecast for the last two years of the data. None of the forecast (except one time in U.S. production) lies outside their individual confidence bars, and thus constancy of the model can be accepted, which is confirmed by the Chow test of stability. Finally, the coefficients of the error correction terms are significant.

Table 5.15. Results of the structural PVEC model (dataset II)

Variable		EXP		SUP		P	
		Coefficient	*t*-value	Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.81	-4.69*	-0.34	-3.08*	-0.96	-2.94*
	EXP_{t-2}			-0.34	-2.95*		
	P_{t-2}					-0.75	-2.59*
	SUP_{t-1}	0.40	2.36*				
Exogenous	HS	0.17	2.14*	0.26	3.22*	0.22	2.73*
	LP	0.12	3.00*	-0.01	-1.24	-0.03	-1.26
	EX	0.28	3.02*	-0.41	-3.49*	-1.00	-3.19*
	MOU (1986:4-1991:4)	-0.02	-0.41	0.06	0.53	-0.03	-0.91
	SLA (1996:2-2001:1)	-0.05	-1.41	-0.01	-0.63	-0.03	-0.81
	SEA_1 (2^{nd} quarter)	0.01	1.09	0.06	2.72*	0.03	1.52
	SEA_2 (3^{rd} quarter)	0.07	1.58	-0.01	-0.36	-0.10	-2.03*
	SEA_3 (4^{th} quarter)	0.09	2.68*	-0.01	-0.17	-0.08	-1.19
Error-correction term		-0.56	-3.64*	-0.46	-2.70*	-0.43	-2.33*
Multivariate tests		$F_{ar}(18,76) = 0.97$; $F_{het}(60,104) = 0.53$; $\chi^2(6) = 11.46$					
LR test of over-identification		$\chi^2(6) = 6.33$					

Note:

1. * indicates significance at the 5% level.
2. Multivariate tests for serial correlation, heteroskedasticity, and normality were not rejected at the 5% significant level.

Figure 5.4. Impulse responses for SPVEC model (dataset II)

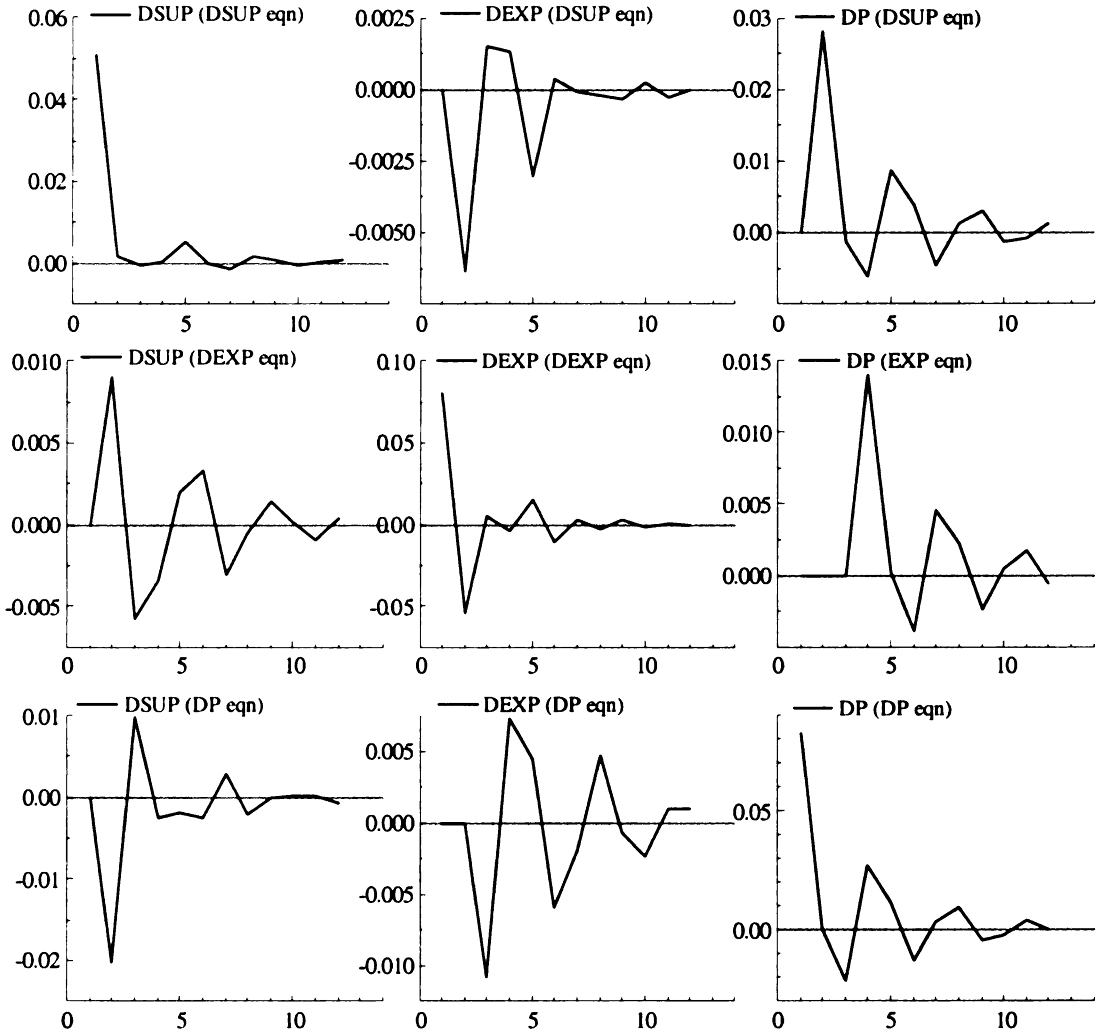
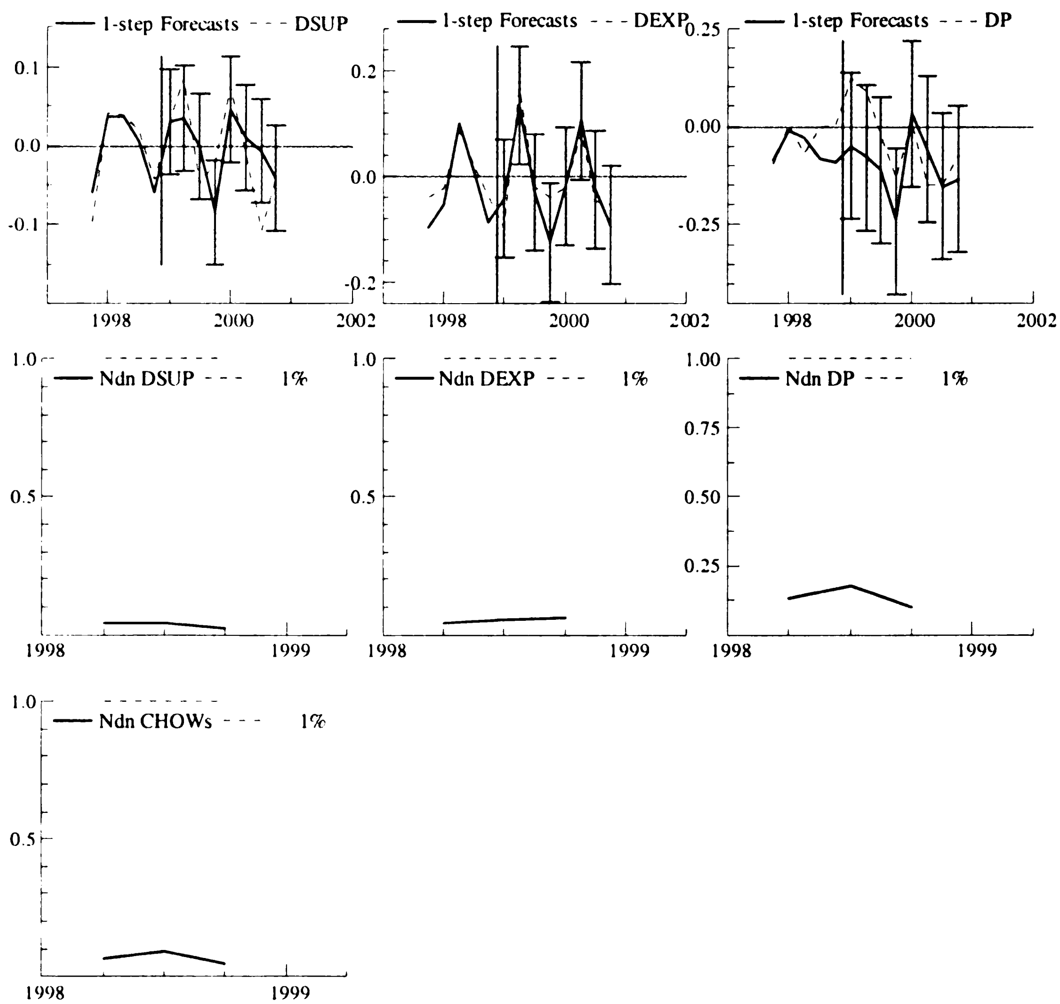


Figure 5.5. One-step model-based forecasts and Chow tests for SPVEC model (dataset II)



5.4. Summary and Conclusions

With two datasets, this study employed the Johansen co-integration test and the VEC model to analyze the U.S. lumber market structure and to estimate the impacts of U.S. restrictions. Although these multiple procedures for the nonstructural modeling were time and effort consuming, they have generated some interesting and consistent results with the structural models. The derived results are thus very helpful for us to understand the U.S.-Canada softwood lumber trade.

The results show that the MOU caused Canada's exports to drop by about 9% and thus pushed up domestic production by about 3%. These results were very close to the findings of the SSE model in which the impact of the MOU was an 8% reduction of Canadian exports. As in the SSE model, however, the SLA had little impact on the lumber market. On the other hand, as a log market shock, the federal harvest reductions in the PNW caused the domestic production to decrease by 4%, which is 1.4% lower than that derived from the SSE model. However, the Asian financial crisis had little effects.

This study also found that lumber prices have little effect on Canadian exports in the short-run. In contrast, other market factors, such as housing starts, exchange rates, and log prices, have significant impacts on Canadian export supply. These results confirm the findings of previous studies (Jennings et al. 1991 and Sarker 1996) that housing starts have a significant effect on the level of Canadian exports. Unlike these two studies, however, this study found that the exchange rate is one of the most significant factors on softwood lumber imported from Canada. Additionally, U.S. log prices affect Canadian supply significantly, which in turn indicates that U.S. producers are sensitive to log prices in their lumber manufacturing. It should be noted that the insignificant impact of lumber prices on Canadian export supply in the short-run does not mean that prices are not an important determinant of export supply in the long run. That is, the results from the co-integration procedure show that the long run response of Canadian exports to prices is significant, which is 0.27 to 0.31. If we interpret the long run response as the upper bound of the price elasticity (Quantitative Micro Software 1998), the price elasticity of Canadian export supply cannot be larger than 0.31 at best. In other words, when lumber price increases by 1%, Canadian export increases by 0.31% at maximum, holding other factors fixed. The price elasticity is larger than that obtained from the SSE model (0.13-0.15), but still much smaller than that derived from previous studies.

CHAPTER 6

WELFARE ESTIMATION OF U.S. RESTRICTIONS

This chapter provides the *ex post* welfare evaluations of the trade restrictions and simulation for the potential effects of the past and proposed future restriction regimes. First, using the price elasticities and the impacts of the trade restrictions estimated from our models, this study evaluates the welfare effects of the MOU and the SLA. Note that this study consistently found that the SLA did not have any significant effects on the lumber market during 1996:2-2001:1. To compare with the results of the previous study (Zhang 2001), however, this study assumes that the SLA significantly reduced Canadian exports by 4%, which is derived from the estimated coefficients in our models. Second, this study simulates the potential impacts of the past and proposed future restriction regimes. The main task of the simulation is to explore more feasible and effective policy options to restrict Canadian lumber exported to the U.S. If the dispute and restrictions are likely to continue in the future, then it is necessary to find which type of restriction — an import tariff, an export tax, or a quota — would be the most effective to regulate lumber imported from Canadian. In so doing, this study will not only provide some estimates of what gains or losses to both countries are under different trade restrictions, but also examine the practicality of the restriction options. This chapter is organized as follows. First, the theoretical framework for the welfare analysis of the trade restrictions is discussed. Next, the market and welfare impacts of the MOU and the SLA are presented. Then, the simulated results for various policy options are discussed. Finally, this chapter ends with summary and conclusions.

6.1. Theoretical Framework

6.1.1. Export Tax

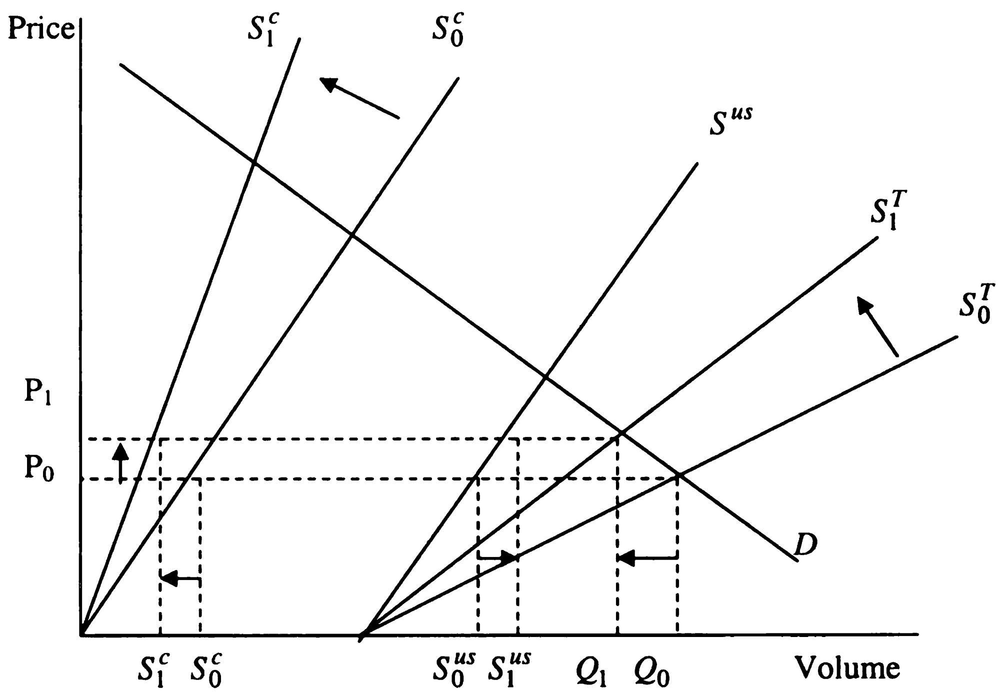


Figure 6.1. Structure of the U.S. lumber market under an export tax

Figure 6.1 illustrates the approach to estimating the market and welfare impacts of an export tax, such as the MOU. Figure 6.1 describes the structure of the U.S. softwood lumber market with a demand function for lumber (D) and a two-part lumber supply with total supply (S^T) equal to the sum of Canadian supply (S^c) and domestic supply (S^{us}). In this model, an *ad valorem* export tax on Canadian lumber results in an inward shift in Canadian export supply (ΔS^c), and a consequent shift in total U.S. supply (ΔS^T). The difference, $\Delta S^T - \Delta S^c$, is then described by movement along the domestic supply curve (ΔS^{us}). An export tax raises the import price at any volume by the tax fraction ($\Delta P = P_1 - P_0$). Hence, the result at the new market equilibrium is a reduction in the imported volume, an increase in U.S. production, an increase in lumber prices, and a reduction in the price received by Canadian producers. Note that the effectiveness of an export tax mainly depends on the price elasticities of demand, domestic supply, and Canadian export supply. For example, if Canadian export supply is sensitive to price changes (elastic), an export tax will produce larger volume reductions and larger price increases in the U.S. than if it were less sensitive to price (inelastic).

The market impacts under an export tax can be calculated once the price elasticities and the estimated impact of an export tax on Canadian lumber are provided. That is, from the definition of elasticity, the changes in total demand, domestic supply, and Canadian export can be calculated as:

$$\Delta Q = \eta \frac{Q_0}{P_0} \Delta P \tag{6.1}$$

$$\Delta S^{us} = \delta \frac{S_0^{us}}{P_0} \Delta P \tag{6.2}$$

$$\Delta S^{c} = \delta^{c} \frac{S_0^{c}}{P_0} \Delta P \tag{6.3}$$

where η, δ and δ^c are the own-price elasticities of demand, domestic supply and Canadian export, respectively. The subscript 0 refers to base year observations. The identity in volume can be defined as $\Delta Q = \Delta S^T = \Delta S^{us} + \Delta S^c$.

The welfare impacts can be directly estimated from the market impacts and Canadian and domestic supply equations as well as the demand equations.[26] Consumer surplus is defined as the area under the demand curve and above the price line. Change in consumer surplus is defined by the total quantity and price changes shown in Figure 6.1:

$$\Delta CS = -[(P_1 - P_0)Q_1 + \frac{1}{2}(Q_0 - Q_1)(P_1 - P_0)] \tag{6.4}$$

Producer surplus is defined as the area above the supply curve and under the price line. Change in domestic producer surplus also follows directly from Figure 6.1:

$$\Delta PS^{us} = (P_1 - P_0)S_0^{us} + \frac{1}{2}(S_1^{us} - S_0^{us})(P_1 - P_0)] \tag{6.5}$$

The change in producer surplus for Canada under an export tax is based on treating the shifted Canadian export curve as the effective supply. Accordingly, change in producer surplus is defined as follows:

$$\Delta PS^c = (1-\alpha)P_1 S_1^c - \frac{1}{2}P_1^* S_1^c - \frac{1}{2}P_0 S_0^c \tag{6.6}$$

$$P_1^* = \frac{P_0}{S_1^c}$$

where α is the export tax rate and P_1^* is the price defined by the pre-policy supply curve at S_1^c. The export tax revenue is defined as $R = \alpha P_1 S_1^c$.

---

[26] These equations are obtained from the price elasticities and base year observations.

(1) Demand: $P = a - \alpha Q$, $\alpha = \frac{P_0}{\eta Q_0}$, $a = P_0 + \alpha Q_0$

(2) Domestic supply: $P = b + \beta S^{us}$, $\beta = \frac{P_0}{\delta S^{us}}$, $b = P_0 - \beta S^{us}$

(3) Canadian export supply: $P = b^c + \beta^c S^c$, $\beta^c = \frac{P_0}{\delta^c S^c}$, $b^c = P_0 - \beta^c S^c$

6.1.2. Tariff-Regulated Quota

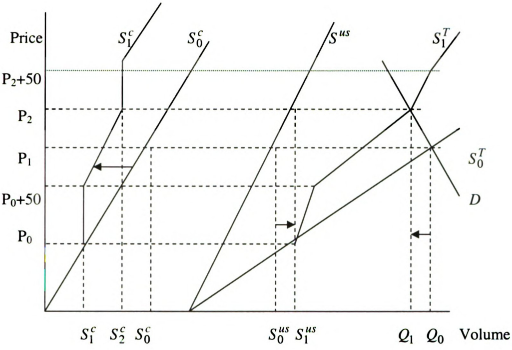


Figure 6.2. Structure of the U.S. lumber market under a tariff-regulated quota

Figure 6.2 demonstrates the method to estimating the market and welfare impacts of a tariff-regulated quota, such as the SLA. As shown in Figure 6.1, the structure of the U.S. lumber market consists of a demand function for lumber (D) and a two-part lumber supply with total supply (S^T) equal to the sum of Canadian export supply (S^c)[27] and domestic supply (S^{us}). Notice that under the SLA, an export fee of \$50/MBF is placed when the total export from the four provinces exceeds 14.7 BBF (S_1^c) until the additional

[27] The SLA imposes an export quota on shipments of softwood lumber from British Columbia, Alberta, Ontario, and Quebec to the U.S. Hence, Canadian export supply in Figure 6.2 includes these four Canadian provinces. In addition, supply from other Canadian provinces is included in domestic supply.

650 MMBF annual limit (ΔS_1^c) is reached. Any additional export beyond S_2^c ($= S_1^c + \Delta S_1^c$) is subject to a fee of \$100/MBF. Hence, the Canadian export supply (S_1^c) and total U.S. supply (S^T) under the SLA become kinked curves.

As seen in an export tax, the market and welfare impacts of the SLA also can be calculated using equation (6.1)-(6.6) once the price elasticities and the estimated quantity impact are provided. Note that there are two differences between the MOU and the SLA to estimating the welfare impacts. First, the price (P_0) when Canadian export supply equal S_1^c can be calculated as:

$$\Delta P_0 = P_2 - \frac{(S_0^c - S_1^c)P_2}{\delta^c S_0^c} \tag{6.7}$$

where δ^C is the price elasticity of Canadian export supply.

Second, change in Canadian producer surplus under the SLA is calculate as:

$$\Delta PS^c = 50S_1^c + \frac{(P_2 - P_0 - 50)(S_1^c + S_2^c)}{2} - \frac{(P_1 - P_0)(S_1^c + S_0^c)}{2} \tag{6.8}$$

The first two components are the portion of the producer surplus where price is above P_0 under the SLA. The third component is the portion of the producer surplus where price is above P_0 under the no-SLA. Note that Canadian producers as a whole paid a \$50/MBF fee for export volume exceeding S_1^c in four years (1997:2-2001:1) and that their total exporting volume was just about equal to or slightly greater than S_2^c. As a result, like Zhang (2001), this study only considered the scenario when Canadian producers paid a \$50 fee and when they collectively exported S_2^c amount of softwood lumber to the U.S.

6.2. Welfare Impacts of Trade Restrictions

6.2.1. Welfare Impacts of the MOU

This study calculated the market and welfare impacts of the MOU using the equations 6.1-6.6 and the price elasticities estimated from the models.[28] Specifically, this study found that the MOU reduced Canadian exports by 8-9% during 1986:4-1991:3. Considering levels of lumber consumption during the same period, an 8-9% reduction in Canadian exports would amount to a reduction of approximately 1.1 BBF annually. However, this does not mean that domestic production would have increased by the same amount during 1986:4-1991:3. Domestic production is influenced by the changes in lumber prices and consumption simultaneously (Figure 6.1). As a result, as long as the demand for lumber is downward sloping, 1.1 BBF is the upper bound on any increase in domestic production caused by the MOU. To estimate these changes in production and consumption, the price elasticities estimated from my models were used to equations 6.1-6.3. Additionally, to estimate the impacts of the MOU, these equations were applied to actual production, consumption, and price data during 1986:4-1991:3. In so doing, this study simulated what the market results would have been if the MOU had not been in place. Note that this study simulated two estimated impacts of the MOU. That is, (1) an 8% reduction in Canadian exports estimated from the SSE model (case I) and (2) a 9% reduction in Canadian exports and a 3% increase of domestic supply obtained from the VAR model (case II).

[28] Demand elasticities (η) = –0.08~-0.13; Domestic supply elasticities (δ) = 0.15~0.26; Canadian export supply (δ^c) = 0.13~0.15.

6.2.1.1. Case I: an 8% Reduction in Exports[29]

This study presents detailed estimates of impacts for the first year (1986:4-1987:3, Table 6.1) and then summarizes total impacts for the period of 1986:4-1991:3 (Table 6.2). The results for the first year showed that Canadian exports declined by 1.1 BBF from the otherwise open market solution, and domestic production increased by 642.2 MMBF. The consequent change in lumber consumption showed a 443.3 MMBF reduction with a price increase of $14.2/MBF. With these market impacts, detailed welfare impacts of the 8% decline in Canadian exports were calculated for the first year. U.S. lumber producers gained $502.5 million, whereas U.S. consumers lost $703.1 million. As a result, estimate net loss to the U.S. was $200.6 million. On the other hand, while Canadian producers lost $269.5 million, the provincial governments gained $417.8 million in export tax revenue. Together, net gains to Canada are $148.3 million. The overall impact of the MOU for the first year was a net loss of $52.3 million, due to U.S. consumer costs exceeding gains to U.S. producers and Canada.

The total and the average annual impacts showed that due to the MOU, Canadian lumber exports reduced by 5.2 BBF during 1986:4-1991:3, whereas U.S. production increased by 3.1 BBF (Table 6.2). As a result, with an annual price increase of $12.9/MBF, U.S. consumption reduced by 2.1 BBF during the same period. With these market impacts, detailed welfare impacts of the 8% decline in Canadian exports were calculated for the full course of the MOU. U.S. lumber producers gained $2.2 billion, whereas U.S. consumers lost $3.1 billion. Together, net loss to the U.S. was $884.1

[29] Case I assumes that a significant drop in Canadian export supply due to the MOU leads to a significant price and thus domestic production increase. In addition, to compare our results with Wear and Lee (1993), estimates of the price and welfare impacts were converted into 1982 U.S. dollars.

million. On the other hand, while Canadian producers lost $1.2 billion, the province governments gained $1.8 billion in export tax revenue. Accordingly, net gains to Canada were $616.5 million. As a result, this analysis placed the deadweight loss at about $267.6 million over the five years, because U.S. producer profits and net gains to Canada were more than offset by larger U.S. consumer costs.

Table 6.1. Results of market and welfare impacts for 1986:4-1987:3 under MOU (case I)

		With MOU	Without MOU	Policy impact
Market impact	Consumption (MMBF)	49,237.0	49,388.1	-151.1 (-0.3)
	Production (MMBF)	35,668.6	34,598.5	+1,070.1 (+3.0)
	Canadian exports (MMBF)	13,568.4	14,789.6	-1,085.5 (-8.0)
	Price (U.S.$/MBF)	205.3	191.1	+14.2 (+6.9)
Welfare impact	U.S. consumer surplus	38,877.9	39,581.0	-703.1 (-1.8)
	U.S. producer surplus	6,370.7	5,868.2	+502.5 (+8.56)
	U.S. net impact	45,248.6	45,449.2	-200.6 (-0.4)
	Canadian producer surplus	2,367.6	2,637.1	-269.5 (-10.2)
	Export tax revenue	417.8	0.0	+417.8
	Canadian net impact	2,785.5	2,637.1	+148.3 (+5.6)
	Deadweight loss (U.S.+ Canadian net impact)			-52.3

Note:

1. All values for welfare impact are in million 1982 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-12.

Table 6.2. Results of market and welfare impacts for 1986:4-1991:3 under MOU (case I)

		Total change	Annual change	% change
Market impact	Consumption (MMBF)	-2,129.4	-425.9	-0.9
	Production (MMBF)	+3,077.8	+615.6	+1.8
	Canadian exports (MMBF)	-5,207.2	-1,041.4	-8.0
	Price (U.S.$/MBF)	-	+12.9	+7.0
Welfare impact	U.S. consumer surplus	-3,076.4	-615.3	-1.8
	U.S. producer surplus	+2,192.3	+438.5	+8.6
	U.S. net impact	-884.1	-176.8	-0.4
	Canadian producer surplus	-1,208.5	-241.7	-10.5
	Export tax revenue	+1,825.0	+365.0	-
	Canadian net impact	+616.5	+123.3	+5.3
	Deadweight loss (U.S. + Canadian net impact)	-267.6	-53.5	-

Note:

1. All values for welfare impact are in million 1982 U.S. $.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-12.

Compared with the results of Wear and Lee (1993), these results indicate that the MOU had a smaller welfare impact on the U.S. lumber market. Specifically, Wear and Lee (1993) found that the MOU led to a 5% reduction of Canadian share in the U.S. market. Given this impact, they estimated market and welfare impacts using the following price elasticities: demand elasticities (η) = –0.17, domestic supply elasticities (δ) = 0.40, and Canadian export supply (δ^c) = 1.0. Notice that they assumed δ^c equal to unity for analytical convenience. They found that the MOU reduced Canadian lumber exports by 10.4 BBF and pushed up U.S. production by 7.3 BBF during 1986-1990. The consequent change in lumber consumption was a 3.1 BBF reduction with an annual price increase of $20/MBF. With these market impacts, U.S. lumber producers gained $2.6 billion for the first four years of the MOU, whereas U.S. consumers lost $3.8 billion, resulting in net U.S. loss of $1.2 billion. These higher market and welfare impacts of the MOU may be due to their higher price elasticities, particularly Canadian export supply.

6.2.1.2. Case II: a 9% Reduction in Exports and a 3% Increase in Production

To estimate the market and welfare impacts of a 9% reduction in Canadian exports and a 3% increase of domestic supply under the MOU, the same procedure used above was adopted. Note that the assumption in case I is no longer valid since the MOU had little effect on lumber price. Hence, the following assumptions were introduced for case II. First, the price is fixed for each year of the MOU. For this reason, the MOU results in an inward shift in Canadian export, and consequent shifts in demand and total supply, including domestic supply. As a result, the new equilibrium is only related to changes in market lumber volume. Second, a change in consumption (ΔDEM) due to the MOU is equal to the subtraction of change in domestic supply (ΔSUP) and change in Canadian export (ΔEXP).

Table 6.3 presents detailed estimates of impacts for the first year (1986:4-1987:3) under the MOU. Canadian exports reduced by 1.2 BBF from the otherwise open market solution, and domestic production increased by 1.1 BBF. The consequent change in lumber consumption shows a 151.1 MMBF reduction with no price effect. With these market impacts, detailed welfare impacts of case II were calculated for the first year. U.S. lumber producers gained $191.1 million, whereas U.S. consumers lost $119.3 million. As a result, the net gain to the U.S. was $71.8 million. On the other hand, while Canadian producers lost $491.9 million, the provincial governments gained $417.8 million in export tax revenue. Together, the net loss to Canada was $74.1 million. Compared to case I, the deadweight loss for the first year shrunk to $2.3 million, with smaller U.S. consumer costs.

Table 6.3. Results of market and welfare impacts for 1986:4-1987:3 under MOU (case II)

		With MOU	Without MOU	Policy impact
Market impact	Consumption (MMBF)	49,237.0	49,388.1	-151.1 (-0.3)
	Production (MMBF)	35,668.6	34,598.5	+1,070.1 (+3.0)
	Canadian exports (MMBF)	13,568.4	14,789.6	-1,221.2 (-9.0)
	Price (U.S.$/MBF)	205.3	205.3	0.0
Welfare impact	U.S. consumer surplus	38,877.9	38,997.2	-119.3 (-0.3)
	U.S. producer surplus	6,370.7	6,179.6	+191.1 (+3.1)
	U.S. net impact	45,248.6	45,176.8	+71.8 (+0.2)
	Canadian producer surplus	2,367.6	2,859.6	-491.9 (-17.2)
	Export tax revenue	417.8	0.0	+417.8
	Canadian net impact	2,785.5	2,859.6	-74.1 (-2.6)
	Deadweight loss (U.S. + Canadian net impact)	-	-	-2.3

Note:
1. All values for welfare impact are in million 1982 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-13.

Table 6.4 summarizes the total and the average annual impacts of the MOU. The results show that due to the MOU, Canadian lumber exports reduced by 5.9 BBF during 1986:4-1991:3, whereas U.S. production increased by 5.1 BBF. As a result, U.S.

consumption reduced by 753.8 MMBF without a price increase. With these market impacts, detailed welfare impacts of case II were calculated for the full years of the MOU. U.S. lumber producers received gains anticipated from such policy; profit increased by $829.0 million. However, the costs to U.S. consumers were much smaller, with $546.7 million. Hence, the net gain to the U.S. was $282.4 million. On the other hand, while Canadian producers lost $2.2 billion, the provincial governments gained $1.8 billion in export tax revenue. Accordingly, the net loss to Canada was $366.1 million. The deadweight loss of the MOU for the five years was a net loss of $83.8 million, which was much smaller than that in case I.

Table 6.4. Results of market and welfare impacts for 1986:4-1991:3 under MOU (case II)

		Total change	Annual change	% change
Market impact	Consumption (MMBF)	-753.8	-150.8	-0.3
	Production (MMBF)	+5,104.2	+1,020.8	+3.0
	Canadian exports (MMBF)	-5,858.1	-1,171.6	-9.0
	Price (U.S.$/MBF)	-	0.0	0.0
Welfare impact	U.S. consumer surplus	-546.7	-109.3	-0.3
	U.S. producer surplus	+829.0	+165.8	+3.1
	U.S. net impact	+282.4	+56.5	+0.1
	Canadian producer surplus	-2,191.2	-438.2	-17.5
	Export tax revenue	+1,825.0	+365.0	-
	Canadian net impact	-366.1	-73.2	-2.9
	Deadweight loss (U.S. + Canadian net impact)	-83.8	-16.8	-

Note:

1. All values for welfare impact are in million 1982 U.S. $.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-13.

Compared to case I, the results of case II showed different size and distribution of welfare gains and losses. Under case I, the MOU placed the maximum burden on U.S. consumers with Canadian producers incurring the second largest loss. Additionally, with consumer costs exceeding producer gains, net impact on the U.S. was negative. In contrast, the Canadian net impact was positive, with export tax revenue exceeding Canadian producer losses. Under case II, however, the ordering of the losses changes: due to little price effects, Canadian producers were the largest losers, with U.S. consumer bearing the second largest burden. In addition, with higher increase in domestic production, U.S. producer gains exceeded consumer losses, which resulted in positive net impact on the U.S. In contrast, Canada suffered a loss in economic surplus, with much larger producer costs exceeding tax revenue.

Furthermore, the results of case II imply that even the SSE model (case I) may be exaggerated the market and welfare impacts of the MOU, due to its assumption that a significant impact in Canadian exports due to the MOU leads to a significant price and thus domestic production increase. However, as seen in case II, the assumption was not valid. That is, the MOU led to the reduction in Canadian exports and thus increase in domestic production, but it did not increase lumber prices. As a result, the deadweight loss under case II was much smaller than that obtained from case I.

6.2.2. Welfare Impacts of the SLA[30]

This study calculated the market and welfare impacts of the SLA using the equations 6.1-6.8 (except 6.6) and the price elasticities estimated from our models. Note

[30] To compare my results with Zhang (2001), estimates of the price and welfare impacts were converted into 1997 U.S. dollars.

that this study consistently found that the SLA did not have any significant effect on the lumber market during 1996:2-2001:1. However, in order to compare with the results of Zhang (2001), this study estimated the market and welfare impacts of the SLA, assuming that the restriction significantly reduced Canadian exports by 4%, which is derived from the estimated coefficients in the models.

This study presents detailed estimates of impacts for the first year (1996:2-1997:1, Table 6.5) and then summarizes total impacts for the period of 1996:2-2001:1 (Table 6.6). The results for the first year showed that Canadian lumber exports reduced by 628.8 MMBF from the otherwise open market solution, and domestic production increased by 354.9 MMBF. The consequent change in lumber consumption showed a 273.9 MMBF reduction with a price increase of $18.3/MBF. With these market impacts, detailed welfare impacts of the 4% decline in Canadian exports were calculated for the first year. U.S. lumber producers gained $585.9 million, whereas U.S. consumers lost $912.1 million. As a result, estimated net loss to the U.S. was $326.1 million. On the other hand, Canadian producers in four provinces gained $237.1 million. It means that even though Canadian producers should reduce their export volume due to the SLA, they gained even more with an increase of lumber price. Additionally, Canadian producers in non-quota constrained provinces gained $39.4 million due to increased exports of lumber to the U.S. Thus, the net gains to Canada were about $327.8 million. However, the overall impact of the SLA for the first year was a net loss of $75.7 million, due to larger U.S. consumer costs.

Table 6.5. Results of market and welfare impacts for 1996:2-1997:1 under SLA

		With SLA	Without SLA	Policy impact
Market impact	Consumption (MMBF)	49,651.0	49,924.9	-273.9 (-0.6)
	Production (MMBF)	32,165.2	31,810.3	+354.9 (+1.1)
	Canadian exports (MMBF)	15,719.0	16,347.8	-628.8 (-4.0)
	Other exports (MMBF)	1,766.8	1,654.3	+112.5 (+6.4)
	Price (U.S.$/MBF)	431.7	413.4	+18.3 (+4.2)
Welfare impact	U.S. consumer surplus	82,442.0	83,354.1	-912.1 (-1.1)
	U.S. producer surplus	12,080.9	11,494.9	+585.9 (+5.1)
	U.S. net impact	94,522.9	94,849.0	-326.1 (-0.3)
	Canadian producer surplus	1,563.4	1,326.4	+237.1 (+17.9)
	Other producer surplus	381.4	341.9	+39.4 (11.5)
	Canadian net impact	2,017.4	1,689.6	+327.8 (+19.4)
	Deadweight loss (U.S. + Canadian net impact)	-	-	-75.7

Note:

1. All values for welfare impact are in million 1997 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-14.
3. Other exports (producer surplus) are the volume (gain) of Canadian exporters in non-quota constrained provinces, such as New Brunswick and Nova Scotia.

The total and the average annual impacts showed that due to the SLA, Canadian lumber exports reduced by 3.1 BBF during 1996:2-2001:1, whereas U.S. production increased by 1.8 BBF (Table 6.6). As a result, with an annual price increase of $14.1/MBF, U.S. consumption reduced by 1.3 BBF during the same period. With these market impacts, detailed welfare impacts of the 4% decline in Canadian exports were calculated for the full course of the SLA. U.S. lumber producers gained $2.5 billion, whereas U.S. consumers lost $3.7 billion. Together, net loss to the U.S. was $1.2 billion. On the other hand, Canadian producers in four provinces and other producers gained

$791.4 million and $200.9 million, respectively. Accordingly, net gains to Canada were $922.3 million. Overall, this analysis places the deadweight loss at about $160.6 million over the five years, because U.S. producer profits and net gains to Canada were more than offset by much larger U.S. consumer losses.

Table 6.6. Results of market and welfare impacts for 1996:2-2001:1 under SLA

		Total change	Annual change	% change
Market impact	Consumption (MMBF)	-1,298.2	-259.6	-0.5
	Production (MMBF)	1,804.8	+361.0	+1.0
	Canadian exports (MMBF)	3,103.0	-620.6	-4.0
	Other exports (MMBF)	+681.5	+136.3	+5.7
	Price (U.S.$/MBF)	-	+14.1	+3.8
Welfare impact	U.S. consumer surplus	-3,689.5	-737.9	-1.0
	U.S. producer surplus	+2,536.6	+507.3	+4.6
	U.S. net impact	-1,152.9	-230.6	-0.3
	Canadian producer surplus	+791.4	+158.3	+7.3
	Other producer surplus	+200.9	+40.2	+2.3
	Canadian net impact	+922.3	+198.5	+7.8
	Deadweight loss (U.S. + Canadian net impact)	-160.6	-32.1	-75.7

Note:

1. All values for welfare impact are in million 1997 U.S. $.
2. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports: 0.13); Welfare results calculated by price elasticities estimated from dataset II are reported in Table A-14.
3. Other exports (producer surplus) are the volume (gain) of Canadian exporters in non-quota constrained provinces, such as New Brunswick and Nova Scotia.

Compared with the results of Zhang (2001), these results indicate that even if the SLA had a significant effect, it would have had much smaller welfare impact on the U.S. lumber market. Specifically, Zhang (2001) found that under the SLA, the estimated annual change in lumber price was approximately $59.1/MBF, or 16%, on average, for the period 1996-2000. Given this price impact, he estimated market and welfare impacts using the following price elasticities: demand elasticities (η) = –0.17, domestic supply elasticities (δ) = 0.40, and Canadian export supply (δ^c) = 0.9. He found that Canadian lumber exports reduced by 12.8 BBF for the first four years under the SLA, whereas U.S. production increased by 7.4 BBF, which in turn decreased lumber consumption of 4.8 BBF. With these market impacts, U.S. lumber producers gained $7.7 billion for the first four years of the SLA, whereas U.S. consumers lost $12.5 billion, resulting in net U.S. loss of $4.7 billion.

These higher welfare impacts of the SLA may be due to the weaknesses of modeling approach (i.e. RFE) and variable use for estimating the restriction effects as well as the adoption of high price elasticities (i.e. Canadian export supply). Particularly, compared to the SSE and VAR findings, the RFE models tend to result in greater restriction effects. In addition, with the problem of excluding a relevant variable (omitted variable bias), Zhang (2001) has overestimated the effect of the SLA (see section 4.3.2.3). Further, even the RFE models could not detect that the SLA had caused the lumber prices to increase by $60/MBF. In addition, even with such a substantial price increase, the welfare gains and losses could not reach billions of dollars a year, because the SSE model showed that the price elasticities in general and that of Canadian export supply in particular are very small.

6.3. Simulation for Trade Policies

The main task of the simulation is to find the more feasible and effective policy options to regulate Canadian lumber exported to the U.S. For instance, if the dispute is likely to continue in the future, then it is necessary to find which type of restriction — tariff, export tax, or quota — would be more effective to restrict Canadian lumber. In so doing, this study can provide some estimates of gains or losses under trade policies to the both countries, and look into the practicality of the options. Like van Kooten (2002), therefore, this study used a partial equilibrium model to analyze the impacts of various trade policy options. For this purpose, this study continues to employ the method of comparing the results of a simulation with no trade restriction (free trade) with a second simulation with a restriction in place.

6.3.1. Scenarios

To analyze the effects of various restriction regimes, realistic free trade volume and price were first determined as the baseline. As noted above, this study simulated what the market results would have been if the SLA had not been in place. In this context, the derived volume and price with no SLA in place can be interpreted as the market condition of unrestricted free trade during the recent periods. Using the average levels of price, demand, and supply including Canadian exports during 1996-2001, the free trade equilibrium volume and price were calibrated as follows: (a) U.S. consumption: 56 BBF,

(b) domestic production: 37 BBF, (c) Canadian exports: 19 BBF, and (d) lumber prices: $400/MBF.[31]

Then, the results of free trade were compared with the simulation of the following three scenarios. That is, (a) an *ad valorem* countervailing duty (CVD) or import tariff with 6.5% and 27.2%, (b) an *ad valorem* export tax with 15% and 20%, and (c) an export quota with the restriction of Canadian exports to 31.5% of U.S. total consumption. This scenario consists of two cases. Quota (I) assumes that Canadian exports only come from four major provinces, whereas quota (II) includes the lumber exports from non-quota constrained provinces. Note that specific tax rates and a quota regulation were chosen based on the past and current trade policies. Chronologically, a 15% export tax was levied by the MOU during 1986:4-1991:3. A 6.5% CVD was then implemented by the U.S. following the termination of the MOU. The U.S. government imposed a 27.2% CVD on Canadian lumber imports with an expiration of the SLA in the early 2001. Most recently, the U.S. and Canada reached a tentative agreement to restrict Canadian exports to 31.5% of U.S. total consumption. Note that a 20% export tax was chosen to compare the results with a 27.2% CVD.

6.3.2. Simulation Results

The estimated results of alternative scenarios indicate that an import tariff has a similar impact on U.S. consumers and producers as an export tax (Table 6.7). That is, with an import tax or export tax, U.S. consumers and Canadian producers are worse off

[31] van Kooten (2002) calculated the free trade volume and price as follows: (a) U.S. consumption: 59 BBF, (b) domestic production: 39 BBF, (c) Canadian exports: 20 BBF, (d) lumber prices: $453/MBF. He calibrated the free-trade equilibrium using the assumed autarkic prices and volume in the U.S. and Canada.

than under free trade, whereas U.S. producers are better off. The difference between the two policies is associated with the distribution of tax revenue. For example, under a 27.2% import tax, the U.S. gain in tariff revenue is \$2.1 billion/year, which more than offsets the price impact of the import tax — the excess of consumer loss over producer gain. As a result, the net U.S. welfare gain is \$1.9 billion/year. This result occurs from the monopsony position of the U.S. in the lumber market and capacity to leverage the terms of trade (Boyd 1988). On the other hand, under a 20% export tax, the tax revenue is sufficient to offset the net Canadian producer loss. As a result, the net gain in Canada from its monopoly position in the lumber market is about \$511.2 million/year. Notice that a 27.2% has a similar impact as a 20% export tax.

A quota scheme basically has a similar impact on U.S. consumers and producers as an import tariff and an export tax; while U.S. consumers are worse off than under free trade, U.S. producers are better off. However, under an export quota, Canadian producers could gain compared to free trade. For example, with the quota scheme, lumber producers in four Canadian provinces could gain \$318 million/year compared to what they would earn under free trade. Additionally, for this simulation, U.S. producers would gain substantially from the quota scheme (about \$600 million/year), which is larger than under a tariff system. Notice that despite the larger size of deadweight loss, the quota scheme is more effective on regulating Canadian exports than other schemes. That is, the quota leads to about a 9% (about 1.7 BBF) average annual reduction in imports from Canada.

Table 6.7. Results of market and welfare impacts under various trade policies

	Free trade	6.5 % import tax	27.2% import tax	15% export tax	20% export tax	Quota (I)	Quota (II)
Change in consumption (MMBF)	56,000	-57.2 (-0.1)	-204.1 (-0.4)	-142.0 (-0.2)	-191.3 (-0.3)	-287.8 (-0.5)	-99.3 (-0.2)
Change in production (MMBF)	37,000	+75.6 (+0.2)	+269.7 (+0.7)	+187.6 (+0.5)	+252.8 (0.7)	+387.2 (+1.1)	+131.2 (+0.4)
Change in Canadian exports (MMBF)	19,000	-132.8 (-0.7)	-473.7 (-2.5)	-329.6 (-1.7)	-444.1 (-2.4)	-1,675.0 (-8.8)	-230.5 (-1.2)
Change in price ($/MBF)	400	+3.1 (+0.8)	+11.2 (+2.8)	+7.8 (+2.0)	+11.0 (+2.7)	+16.1 (+4.0)	+5.5 (+1.4)
U.S. consumer surplus ($ million)	+86,153	-175.9 (-0.2)	-626.8 (-0.7)	-436.3 (-0.5)	-587.6 (-0.6)	-883.2 (-1.1)	-305.2 (-0.4)
U.S. producer surplus ($ million)	+12,876	+116.4 (+0.9)	+416.4 (+3.2)	+289.3 (+2.3)	+390.3 (+3.0)	+598.8 (+4.7)	+202.2 (+1.6)
Government revenue ($ million)	-	+495.2	+2,072.2	-	-	-	-
Net U.S. impact ($ million)	+99,029	-59.5 (+420.8)	-210.4 (+1,861.8)	-147.0	-197.4	-284.4	-103.0
Canadian producer surplus ($ million)	+3,800	-229.4 (-6.1)	-805.4 (-21.2)	-634.5 (-16.7)	-818.4 (-21.5)	+318.3 (+8.4)	+101.6 (+1.4)
Government revenue ($ million)	-	-	-	+1,142.1	+1,529.6		
Net Canadian impact ($ million)	+3,800	-229.4	-805.4	+507.6	+511.2	+318.3	+101.6

Note:

1. Quota (I) restricts lumber export from four Canadian provinces to 31.5% of U.S. total consumption, while quota II includes exports to the U.S. by non-quota constrained provinces.
2. Values in parentheses are percentage changes, which is relative to free trade; net U.S. impact values in parentheses are total U.S. gains with tariff revenue.
3. Price elasticities estimated from dataset I are adopted (demand: -0.13, domestic supply: 0.26, and Canadian exports.

However, as van Kooten (2002) pointed out, the effectiveness of a quota on restricting Canadian lumber depends on whether or not all Canadian provinces are included in the system. For example, while the quota scheme I restricts Canadian exports from four Canadian provinces, lumber exports from non-quota constrained provinces undermine the effectiveness of a quota. The quota II shows that if other Canadian provinces other than the four provinces are excluded under a quota scheme, the policy leads to only a 1.2% (about 230.5 MMBF) average annual reduction in Canadian exports. Therefore, it is crucial to include all Canadian provinces under a quota scheme to restrict Canadian exports effectively.

In addition, the effectiveness of a restrictive regime depends on the assumption about the price elasticities of demand, domestic supply, and Canadian export supply. Particularly, changes in the price elasticity of Canadian export supply alone significantly alter the effectiveness of a restrictive regime (Boyd and Krutilla 1987). Hence, this study simulated impacts of an import tariff and an export tax using different price elasticities of Canadian export supply (δ^c). Note that while $\delta^c = 0.31$ is estimated from our models, $\delta^c = 0.9$ is obtained from previous studies (i.e. Zhang 2001). The effects of changes in the price elasticity of Canadian export are summarized in Table 6.8. The results indicate that higher price elasticity of Canadian export leads to a significant reduction of lumber imported from Canada. For example, with δ^c=0.9, a 27.2 % import tariff pushes Canadian exports down by 10.7% (about 2.0 BBF) annually, which is much higher than under a quota scheme. The same holds true under a 20% export tax. Therefore, the results imply that the price elasticities play crucial role in deciding the effectiveness of a restrictive regime.

Table 6.8. Estimated economic impacts of import and export tax

	Elasticity	27.2% import tax		20% export tax		15% export tax	
		Change	%	Change	%	Change	%
Canadian exports (MMBF)	0.13	-437.7	-2.5	-444.1	-2.4	-329.6	-1.7
	0.31	-988.6	-5.2	-925.1	-4.9	-681.6	-3.6
	0.90	-2,036.6	-10.7	-1,898.5	-10.0	-1,379.0	-7.3
Price ($/MBF)	0.13	+11.2	+2.8	+11.0	+2.7	+7.8	+2.0
	0.31	+23.4	+5.9	+21.9	+5.5	+16.1	+4.0
	0.90	+48.2	+12.1	+44.9	+11.2	+32.6	+8.2
U.S. consumer surplus ($ million)	0.13	-626.8	-0.7	-587.6	-0.6	-436.3	-0.5
	0.31	-1,305.4	-1.5	-1,221.7	-1.4	-901.0	-1.1
	0.90	-2,678.2	-3.1	-2,497.9	-2.9	-1,818.1	-2.1
U.S. producer surplus ($ million)	0.13	+416.4	+3.2	+390.3	+3.0	+289.3	+2.3
	0.31	+872.4	+6.8	+815.9	+6.3	+600.0	+4.7
	0.90	+1,811.5	+14.1	+1,686.8	+13.1	+1,220.4	+9.5
Government revenue ($ million)	0.13	+2,072.2	-	-	-	-	-
	0.13	+2,074.3	-	-	-	-	-
	0.90	+2,068.0	-	-	-	-	-
Net U.S. impact ($ million)	0.13	+1,861.8	+1.9	-197.4		-146.9	-0.2
	0.31	+1,641.3	+1.7	-405.9	-0.4	-300.0	-0.3
	0.90	+1,201.3	+1.2	-811.1	-0.8	-597.6	-0.6

6.4. Summary and Conclusions

This study evaluated the *ex post* welfare consequences of the trade restrictions and simulated the potential effects of the past and proposed future restriction regimes. The results of the welfare analyses reflect effective rent seeking by domestic producers and/or Canadian producers. For example, the MOU caused Canadian exports to drop by 8-9% during 1986:4-1991:3, which in turn led to considerable improvement in the competitive position of U.S. producers in the domestic market. As a result, under the MOU, U.S. producers increased domestic production by 1.8-3.0% and gained more than $615 million/year (in 1982 dollars). In addition, under the SLA, despite a 4% reduction of their exports, Canadian producers and the Canadian provincial government gained about $350 million/year (in 1997 dollars) in the U.S. market. In contrast, U.S. consumers lost about $110-$615 million/year (in 1982 dollars) under the MOU and about $740 million/year (in 1997 dollars) under the SLA, respectively. As a result, the trade restrictions can be seen as effective measures of welfare transfer from U.S. consumers to U.S. producers and/or Canadian producers.

However, this study showed that compared with the results of Wear and Lee (1993) and Zhang (2001), the MOU and the SLA had moderate welfare impacts on the U.S. lumber market. It is because in addition to little price effect estimated by the VAR model, the SSE model found that the price elasticities in general and that of Canadian export supply in particular are very small. The results of this study thus showed that the net loss to the U.S. under the MOU was at maximum $884 million with an estimated price increase of $12.9/MBF (in 1982 dollars) during 1986:4-1991:3. Additionally, the net loss to the U.S. under the SLA was about $1.2 billion with an estimated price impact of $14.1/MBF (in 1997 dollars) during 1996:2-2001:1. In contrast, the results of Wear and Lee (1993) showed that an estimated price increase of $20/MBF (in 1982 dollars) resulted in about $1.2 billion of the net U.S. loss. Furthermore, Zhang found that the SLA caused lumber prices to go up by approximately $60/MBF (in 1997 dollars), which in turn resulted in $4.7 billion of the net U.S. loss for the first 4 years under the SLA (1996-2000). Hence, the results of the welfare consequences confirmed the two findings in previous chapters. First, compared to the SSE and VAR findings, the RFE models tend to overestimate the effects of the trade restrictions. Second, considering the crucial role that price elasticities play in generating welfare gains and losses, the higher price elasticities used in previous studies may result in significant welfare consequences of the restrictions.

The simulation results demonstrated that from the U.S. perspective, the best strategy for a restrictive regime is the quota scheme that includes lumber producers in all Canadian provinces since it provides the greatest reduction of Canadian exports to the U.S. For example, the quota scheme leads to about a 9% (about 1.7 BBF) average annual reduction of Canadian lumber. Furthermore, U.S. producer benefits could be about $600 million/year, which is larger than under a tariff system. The results also suggested that as a second best alternative, the U.S. could employ an import tariff and directly collect the tax revenue by the trade restriction. As van Kooten (2002) suggested, on the other hand, the best strategy for Canada is also the quota system that includes all Canadian provinces. This scheme provides the greatest benefits (about $320 million/year) to Canadian producers. As a second best alternative, Canada could employ an export tax and directly collect the tax revenue by the trade restriction. Therefore, the quota scheme including all

Canadian provinces could be a mutually acceptable strategy for a restrictive regime because it provides not only the efficient restriction of Canadian exports but also the greatest benefits to both U.S. and Canadian producers, all at the expense of U.S. consumers.

CHAPTER 7

CONCLUSIONS AND POLICY IMPLICATIONS

7.1. Summary of Results

With an expanded list of variables for the period 1980:1-2001:4, this study has developed both the structural SSE and the nonstructural VAR models to analyze the U.S. softwood lumber market structure and to assess the market and welfare impacts of the U.S. restrictions. In addition, this study employed the RFE to detect the potential variations between this and previous studies, and to discover the differences between the RFE results and those obtained from the SSE and VAR models.

The SSE and VAR models showed that the MOU caused Canadian exports to drop by 8-9% during 1986:4-1991:3, which in turn pushed up domestic production by 3.0%. But it did not lead to a significant increase of lumber prices. However, the SLA did not have any significant impact on Canadian exports, domestic production, and lumber prices during 1996:2-2001:1. On the other hand, the RFE models for lumber prices and Canadian export supply resulted in greater MOU impacts. That is, the MOU pushed up lumber prices by 11-13%, and caused the Canadian export supply to drop by as much as 13%. However, even with the RFE models, this study could not detect that the SLA had a significant impact on lumber prices or Canadian export supply (Table 7.1).

The SSE model showed that the price elasticities of demand for and supply of lumber and Canadian export supply are generally very low. Particularly, the price elasticity of Canadian export supply (0.13-0.15) is much smaller than that used by previous studies (0.6-0.9) for welfare analysis. The VAR model corroborates this finding;

lumber prices do not have a significant effect on Canadian exports in the short run, while even in the long run the price response of Canadian exports is cannot be larger than 0.31 at maximum. In contrast, market factors such as housing starts, log prices, and exchange rates, have played key roles in shaping the U.S. lumber market. Particularly, the dynamics of log market have affected domestic and Canadian export supply significantly. For example, the federal harvest reductions in the PNW caused the domestic production to decrease by 4.0-5.4%, which in turn resulted in lumber prices and volume changes (Table 7.1).

Table 7.1. Summary of impacts of the trade restrictions and log market dynamics

	SSE model	VAR model	RFE model
MOU (1986:4-1991:3)	- Significant * EXP: -8%	- Significant * EXP: -9% * SUP: +3% * P: insignificant	- Significant * EXP: -13% * P: +11-13%
SLA (1996:2-2001:1)	- Insignificant	- Insignificant	- Insignificant
Harvest reductions (1991:1-1993:4)	- Significant * SUP: -5.4%	- Significant * SUP: -4%	- Significant * P: +15%
Asian financial crisis (1997:3-1998:4)	- Insignificant	- Insignificant	- Insignificant

Note:
1. EXP, SUP, and P represent Canadian export supply, domestic production, and lumber prices, respectively.

This study found that compared with the results of previous studies, the MOU and the SLA had moderate welfare consequences on the U.S. lumber market. For example, the net loss to the U.S. under the MOU was about $884 million/year (in 1982 dollars) at maximum, with an annual price increase of $12.9/MBF. Wear and Lee (1993) estimated that the MOU caused lumber prices to increase by $20/MBF annually, resulting in about $1.2 billion/year of the net U.S. loss. In addition, even if the SLA is assumed to be effective, the restriction caused lumber prices to go up by $14.1/MBF (in 1997 dollars) annually, which in turn resulted in about $1.2 billion/year of the net U.S. loss. Zhang (2001) estimated that the net loss to the U.S. under the SLA was about $4.7 billion/year (in 1997 dollars), with an annual price increase of about $60/MBF.

Finally, the simulation results demonstrated that from the U.S. perspective, the best strategy for a restrictive regime is the quota scheme covering all Canadian lumber producers, as long as the U.S. ignores the well being of its consumers. This provides the greatest reduction of lumber imported from Canada as well as the greatest benefits to U.S. producers. Alternatively, the U.S. could make use of an import tariff to regulate Canadian exports as well as to collect the tax revenue by the restriction. However, the results suggest that the most worrisome regime would be an export tax. That is, an export tax results in large Canadian government revenues and less benefits to U.S. producers than under a quota or an import tax.

7.2. Conclusions and Policy Implications

This study advances a methodological as well as a policy contribution. The SSE and VAR models provide a substantial contribution for analyzing the U.S. lumber market structure as well as assessing the impacts of the trade restrictions. The SSE model mitigated the RFE's vulnerability to problems of endogeneity, simultaneity, and collinearity, and used all the information from the full system of equations to obtain efficient estimates of the trade restriction effects. Further, the SSE model derived the price elasticities, so that this study was able to contrast the estimates with those of previous studies and to calculate the welfare effects directly. Nonetheless, the SSE model has its shortcomings, mainly because of the absence of a true structure to capture the underlying market and the lack of dynamic interaction of the relevant variables (Hamilton 1994). Moreover, considering the way the SSE model is constructed, it cannot directly capture all the effects of a trade restriction, particularly those on lumber prices, Canadian exports, and U.S. production. Instead, the effects of the restriction on U.S. production and lumber prices are normally inferred by solving the SSE. For example, the SSE model estimated the impact of the MOU with the presence of a dummy variable in the Canadian export supply. For the welfare analysis, this study then solved the SSE simultaneously for the equilibrium values of other endogenous variables — lumber prices and U.S. production. To this end, the SSE procedure assumed that a significant drop in Canadian exports due to the MOU leads to significant price and thus U.S. production increase.

The VAR model was well suited to explore the dynamic linkages of related variables without imposing *a priori* theoretical structure. Moreover, the VAR model demonstrated the ability to capture both the short run and long run responses of an endogenous variable to other variables endogenous to the system, and its revelation of the potential effects of a trade restriction on every endogenous variable. For example, the VAR showed that the MOU caused changes in two endogenous variables — a reduction in Canadian exports and thus an increase in U.S. production, but had little impact on the other endogenous variable — lumber prices. In addition, the VAR found that the impact of lumber prices on Canadian exports is little in the short run and is quite small even in the long run. This implies that even the SSE may exaggerate the market and welfare impacts of the trade restrictions, as long as it assumes that a significant impact in one endogenous variable translates into significant impacts in others. In fact, two welfare analyses of the MOU confirmed this speculation. That is, with little price effect (case II), the market and welfare impacts of the MOU are quite smaller those derived from the SSE (case I). From this perspective, therefore, the nonstructural VAR analysis was well served to identify the relationships of the U.S. lumber market.

However, it should be noted that employing both the SSE and the VAR models have helped this study to draw more balanced and consistent conclusions. First, this approach helped estimate the impacts of the trade restrictions in a corroborative way. For example, the SSE and VAR models generated quite consistent results of the MOU effect — an 8% or a 9% reduction in Canadian exports. This finding thus leads to compare the estimates with those from previous studies and conclude that the RFE tends to overestimate the effects of the trade restrictions. Second, this approach allowed taking advantages of both systems — deriving the price elasticities from the SSE and identifying the market dynamics from the VAR. Thus, these complementary features have greatly enhanced the knowledge of various aspects of the lumber trade and the quality of this study in the hotly contested area.

On the other hand, the results showed that market factors other than lumber prices have significant impacts on consumption, production, and Canadian exports. In other words, it is crucial to understand the influences of other market factors and their dynamic interactions (i.e. log markets) other than the trade restrictions in order to resolve the dispute. In addition, while U.S. log prices affect domestic production and Canadian export supply significantly, Canadian log prices have no significant impact on its exports.

These results might have to do with the different stumpage pricing systems and limited extent of the log market. It might also imply that, unlike their U.S. counterparts, Canadian producers are insensitive of log prices in their lumber manufacturing, due mainly to lower log prices caused by restrictions on log exports. The results further suggest that it is true to a large extent; Canadian exports are driven by U.S. demand-side factors, such as housing starts and mortgage rates. However, Canadian exports have also to do with U.S. supply side factors such as manufacturing capacity and log prices. This study, therefore, concludes that demand as well as supply factors have influenced on shaping the U.S. softwood lumber market.

The welfare analyses showed that the trade restrictions had moderate welfare consequences on the U.S. lumber market, compared with the results of previous studies estimated by the FRE. It is because, other than modest impacts of the restrictions, the SSE shows that the price elasticities in general and that of Canadian export supply in particular are very small. As a result, the welfare analyses confirm that an RFE model must have exaggerated the welfare consequences with the combined use of the overestimated market effects of the restrictions and higher price elasticities, borrowed or assumed.

This study also found that the best strategy for a restrictive regime to regulate Canadian exports is the quota scheme covering all Canadian lumber producers. However, It should be emphasized that, while a quota scheme is the best strategy for a restrictive regime, it should overcome the following potential shortcomings to regulating Canadian lumber effectively. First, in addition to including all Canadian provinces, a quota scheme also should encompass all softwood lumber products. For example, any supplies of noncovered Canadian lumber, such as roughers and predrilled studs under the SLA, could undermine the effectiveness of the scheme by decreasing lumber prices in the market. Second, a quota scheme needs to abandon a trigger-price mechanism, which simply constitute another means by which some Canadian producers can get around the system (van Kooten 2002).

7.3. Future Research

This study has taken macroeconomic approach to analyze the U.S. lumber market structure and thus to assess the market and welfare impacts of the trade restrictions. Given the overall picture of the North American lumber market, future research on the following issues could be conducted. First, after the expiration of the SLA, the Canadian provincial governments, particularly British Columbia (BC) proposed its forest policy change measures to seek a durable solution of the dispute. To this end, BC's proposal provided a framework of stumpage pricing system determined by competitive auctions. Thus, it would be an interesting research area to assess the progress for the policy changes and thus to contrast this with U.S. stumpage pricing system.

Second, this study found that exchange rates have a substantial effect on the level of Canadian lumber exported to the U.S. However, this result puts at odds with other studies that Canadian exports might have had little to do with the exchange rate (Jennings et al. 1991, Sarker 1996). Thus, it would be worthwhile to revisit this issue by assessing the pass-through from exchange rates to import prices. Finally, a study on production efficiency of lumber industry in the U.S. and Canada could provide useful information on the Canadian producers' claim that they have extended their market share in the U.S. through competitive advantage of production efficiency.

LITERATURE CITED

Adams, D.M. 2003. Market and resource impacts of a Canadian lumber tariff. Journal of Forestry 101(1): 48-52.

Adams, D.M. and R.W. Haynes. 1996. The 1993 timber assessment market model: structure, projections and policy simulations. Portland, OR: USDA Forest Service, General Technical Report PNW-GTR-368.

Adams, D.M. and R.W. Hayne. 1985. Changing perspective on the outlook for timber in the U.S. Journal of Forestry 83(1): 32-35.

Adams, D.M., B.A. McCarl and L.H. Homayounfarrokh. 1986. The role of exchange rates in Canadian-U.S. lumber trade. Forest Science, 32: 973-988.

American Forest and Paper Association. Various months. Monthly statistics for the wood product industry. Washington, DC.

American Forest and Paper Association. 2001. U.S. Forest Facts and Figures. Washington, DC.

Anderson, F.J. and R.D. Cairns. 1998. The softwood lumber agreement and resource politics. Canadian Public Policy 14: 186-196.

Arbor-Pacific Forestry Service. The 2000 Log Lines Statistical Yearbook. Mount Vernon, WA.

Banerjee, A., R.L. Lumsdaine and J.H. Stock. 1992. Recursive and sequential tests of the unit root and trend-break hypothesis: theory and international evidence. Journal of Business & Economic Statistics 10: 271-287.

BC STATS. May 2001. Ministry of Finance and Corporate Relations. British Columbia.

Blough, S.R. 1992. The relationship between power and level for generic unit root tests in finite samples. Journal of Applied Econometrics 7: 295-308.

Boyd, R. 1988. The politics and consequences of protectionism: a case study in North American Lumber Market. Journal of Policy Modeling 10(4): 601-609.

Boyd, R. and K. Krutilla. 1987. The welfare implications of U.S. trade restrictions against the Canadian softwood lumber industry: a spatial equilibrium analysis. Canadian Journal of Economics 20: 17-35.

Brander, J. and B. Spencer. 1985. Export subsidies and international market share rivalry. Journal of International Economics 18: 82-100.

Buongiorno, J., J.P. Chavas and J. Uusivouri. 1988. Exchange rates, Canadian lumber imports and U.S. prices: a time series analysis. Canadian Journal of Forest Research 18: 1587-1594.

Buongiorno, J, J.J. Chou and R.N. Stone. 1979. A monthly model of the United States demand for softwood lumber imports. Forest Science 25: 641-655.

Bureau of Economic Analysis. http://www.bea.doc.gov.

Chao, J.C. and P.C.B. Phillips. 1999. Model selection in partially non-stationary vector autoregressive processes with reduced rank structure. Journal of Econometrics 91(2): 227.271.

Cashore, B. 1998. An examination of why a long-term resolution to the Canada-U.S. softwood lumber dispute eludes policy markets. Canadian Forest Service, working paper 98-02.

Cashore, B. 2001. What should Canada do when the softwood lumber agreement expires? Forest Policy Center. Auburn University.

Chen N.J., G.C.W. Ames and A.L. Hammett. 1988. Implications of a tariff on imported Canadian softwood lumber. Canadian Journal of Agricultural Economics 36: 69-81.

Coalition for Fair Lumber Imports. 2000. Canadian government should end lumber subsidies and adopt competitive timber system. Washington, DC.

Coalition for Fair Lumber Imports. 2001. Canada's lumber subsidies are destroying the U.S. lumber industry (http://www.fairlumbercoalition.org).

Cohen D.H. 1996. A review of structural change in North American consumption of softwood lumber: past, present and future. The Forest Chronicle 72(6): 631-636.

Constantino, L.F. and D. Haley. 1989. A comparative analysis of sawmilling productivity on the British Columbia coast and in the U.S. douglas-fir region:1957-1982. Forest Product Journal 39(4): 57-61.

Constantino, L.F. and M.B. Percy. 1991. The political economy of Canada-U.S. trade in forest products. In Uhler (ed): Canada-U.S. trade in Forest Products (pp.52-72). Vancouver, BC: University of British Columbia Press.

Darnell, A. 1994. A Dictionary of Econometrics. Edward Elgar Publishing, UK.

Darr, D.R. 1977. Floating exchange rates and log export policy. Journal of Forestry 75(2): 88-90.

Department of Foreign Affairs and International Trade, Canada 2001. http://www.dfaitmaeci.gc.ca/~eicb/softwood/

Dickey, D.A. and W.A. Fuller. 1971. Distribution of the estimators for auto-regressive time series with a unit root. Journal of the American Statistical Association 74: 427-431.

Doornik, J. and D., Hendry. 1994. Interactive econometric modeling of dynamic system (PcFiml 8.0). University of Oxford Institute of Economics and Statistics. London: International Thomson Publishing.

Doornik, J. and D. Hendry. 2001. PcGive 10. Empirical Econometric Modeling. London, UK: Timberlake Consultants Ltd.

Emerson, D. 2001. Canada-U.S. lumber trade: past mistakes, future lessons (http://www.forestweb.com/editorial/commentary/030701.html).

Engle, R. and C.W.J Granger. 1987. Co-integration and error-correction: representation, estimation and testing. Econometrica 55: 251-276.

FAO. 2001. FAO Statistical Database. http://apps.fao.org.

Feinberg, R.M. 1989. Exchange rate and unfair trade. Review of Economics and Statistics 71(4): 704-707.

Finger, J.M., H.K. Hall and D.R. Nelson. 1982. The political economy of administered protection. American Economic Review 72: 452-466.

Flynn, B. 2001. Impact of exchange rate changes on U.S. forest products. Timber Mart-South Market Newsletter 6(3): 10-12.

Fukuda, J. 2001. A Study of the Effects of the Canada-U.S. Softwood Lumber Agreement. Center for International Trade in Forest Products. University of Washington.

Garcia, J.P., B. Lippke and J. Baker. 1997. Trade barrier in the pacific forest sector: who wins and who losses. Contemporary Economic Policy 15: 87-103.

Gorte, R.W. and J. Grimmett. 1996. Softwood lumber imports: The 1996 U.S.-Canada Agreement. Washington, DC: Congressional Research Service.

Gorte, R.W. and J. Grimmett. 2002. Lumber imports from Canada: issues and events. Washington, DC: Congressional Research Service (updated May 9).

Greene, W. 1993. Econometric Analysis. New York: Macmillan Publishing Co.

Gregory, A.W., J.M. Nason and D. Watt. 1996. Testing for structural breaks in co-integrated relationships. Journal of Econometrics 71: 321-341.

Griffith, S.S. 2001. Certain softwood lumber from Canada: comments for preliminary determination, submitted to Donald L. Evans, Secretary of Commerce on behalf of the Government of British Columbia. Vancouver, BC: Akin, Gump, Strauss, Hauer & Feld LLP.

Haley, D. 1980. A regional comparison of stumpage values in BC and the U.S. Pacific Northwest. Forestry Chronicle 56: 225-250.

Hamilton, J. 1994. Times Series Analysis. Princeton, NJ: Princeton University Press.

Harris, R. 1995. Using Co-integration Analysis in Econometric Modeling. London, UK: Prentice Hall.

Hatemaki, L, R. Hanninen and A. Toppinen. 2001. A System for Short-Term Forecasting of the Finnish Forest Sector (MESU). Finnish Forest Research Institute.

Hayter, R. 1992. International trade relations and regional industrial adjustment: the implications of the 1982-86 Canadian-U.S. softwood lumber dispute for British Columbia. Environment and Planning 24: 153-170.

Hendry, D. 1995. Dynamic Econometrics. Oxford, UK: Oxford University Press.

Hoberg, G. and P. Howe. 1999. Law, knowledge, and national interests in trade disputes: the softwood lumber case. UBC Institute of International Relations, working paper No. 29.

Howard, J.L. 2000. U.S. timber production, trade, consumption, and price statistics 1965-1999. USDA. Research Paper FPL-RP-595 (http://www.fs.fed.us).

International Monetary Fund. Various years. International Financial Statistics. Washington, DC.

Irland, L.C. 1986. Canada-U.S. forest products trade: tensions in a maturing market. Forest Product Journal 37: 21-29.

Irland, L.C. 1995. Perspectives on the changing economic structure of the North American softwood lumber industry. Forest Product Journal 45: 53-65.

Jennings, S., W. Adamowicz and L. Constantino. 1991. The Canadian lumber industry and the macroeconomy: a vector auto-regression analysis. Canadian Journal of Forest Research 21: 288-299.

Johansen, S. 1995. Likelihood-based inference in co-integrated vector auto-regressive models. Oxford, UK: Oxford University Press.

Johansen, S. and K. Juselius. 1990. Maximum likelihood estimation and inference on co-integration — with application to the demand for money. Oxford Bulletin of Economics and Statistics 52: 169-210.

Kaiser, H.F. 1984. Floating exchange rates and the U.S. forest products trade balance. Forest Product Journal 34: 55-58.

Kalt, J.P. 1988. The political economy of protectionism: tariffs and retaliation in the timber industry. In R.E. Baldwin (ed): Trade Policy Issues and Empirical Analysis (pp.339-368). Chicago, IL: University of Chicago Press.

Latta, G.S. and D.M. Adams. 2000. An econometric analysis of output supply and input demand in the Canadian softwood lumber industry. Canadian Journal of Forest Research 30(9): 1419-1928.

Lindsey, B., M.A. Groombridge and P. Loungani. 2000. The economic impact of trade protection of the softwood lumber industry. Washington, DC: CATO.

MacCarl, B.A. and R.W. Hayne. 1985. Exchange rates influence softwood lumber trade. Journal of Forestry 83(1): 368-370.

MacKinnon, M. and P. Kennedy. August 11, 2001. Canadians incensed over duty on lumber. The Globe and Mail (http://www.theglobeandmail.com).

Maddala, G.S. and In-Moo Kim. 1998. Unit roots, co-integration, and structural change. Cambridge, UK: Cambridge University Press.

Martin, R.M. and D.R. Darr. 1997. Market responses to the U.S. timber demand-supply situation of the 1990s: implications for sustainable forest management. Forest Prodcut Journal 47: 27-32.

Mas-colell, A, M.D. Whinston and J.R. Green. 1995. Microeconomic Theory. Oxford, UK: Oxford University Press.

Murray, B.C. and D.N. Wear. 1998. Federal timber policy and interregional arbitrage in U.S. lumber: an event study. Land Economics 74: 76-91.

Myneni, G., J.H. Dorfman and G.C.W. Ames. 1994. Welfare impacts of the Canada-U.S. softwood lumber trade dispute: beggar thy consumer trade policy. Canadian Journal of Agricultural Economics 42: 261-271.

National Association of Home Builders. 2001 (Oct. 31). Lumber ruling harms consumers and hurts fragile economy (http://www.nahb.com/newa/default.htm).

Perron, P. 1989. The great crash, the oil shock and the unit root hypothesis. Econometrica 57: 1361-1402.

Quantitative Micro Software. 1998. EViews User's Guide. Irvine, CA.

Quintos, C.E. 1995. Sustainability of the deficit process with structural shifts. Journal of Business and Economic Statistics 13: 409-417.

Ragosta, J.A. and H.L. Clark. 2000. Canadian governments should end lumber subsidies and adopt competitive timber system, comments submitted to the Office of the United States Trade Representative on Behalf of the Coalition for Fair Lumber Imports. Washington, DC: Dewey Ballantine LLP.

Random Lengths. Various issues. Random Lengths yearbook. Eugene, OR.

Random Lengths. U.S.-Canada trade dispute timeline (http://www. Randomlengths.com/newtimeline.html).

Sarker, R. 1996. Canadian softwood lumber export to the U.S.: a co-integrated and error corrected system. Journal of Forest Economics 2: 205-231.

Sierra Legal Defense Fund. 2001. Stumpage sellout: how forest company abuse of the stumpage system is costing BC taxpayers millions. Vancouver, BC.

Sims, C.A. 1980. Macroeconomics and reality. Econometrica 48:1-48.

Singh, B.K and J.C. Nautiyal. 1986. An econometric analysis of markets for Canadian lumber. Wood and Fiber Science 18(3): 382-396.

Spelter, H. 1992. Technology-driven substitution in the forest sector: the variable price elasticity model revisited. In Lonnstedt, Lars (ed): Forest Sector Analysis: Proceedings of P06.02 FORESEA, IUFRO Centennial. Berlin, Germany (8/4-9/30).

Tougas, F. 1988-89. Softwood lumber from Canada: natural resources and the search for a definition of countervailable domestic subsidy. Gonzaga Law Review 24: 135-165.

Uhler, R. 1991. Canadian public timber pricing and the great subsidy debate. In Uhler (ed): Canada-U.S. Trade in Forest Products (pp. 73-93). Vancouver, BC: University of British Columbia Press.

USDA Forest Service. 2001. 2000 RPA Assessment of forest and range lands. Washington, DC.

USDA Foreign Agricultural Service. 2000. Wood Products: International Trade and Foreign Market (Annual Statistical Trade Issue). Washington, DC.

U.S. Census Bureau. www.census.gov/const/C20/startsua.pdf.

van Kooten, G.C. 2002. Economic analysis of the Canada-United States softwood lumber dispute: playing the quota game. Forest Science 48(4): 712-721.

Wear, D.N. and K.J. Lear. 1993. U.S. policy and Canadian lumber: effects of the 1986 Memorandum of Understanding. Forest Science 39: 799-815.

Wooldridge, J.M. 2000. Introductory Econometrics: a Modern Approach. South-Western College Publishing.

Worden, V.L. 2000. NLBMDA determined to defeat lumber trade restraints (http://fash.lakeadu.ca/~mshannon/softwood2.htm).

Yin, R. 2001. Spotted owl protection, booming housing market, and the timber price movements in the Pacific Northwest. Natural Resource Modeling 14(3): 1-18.

Yin, R and J. Baek. 2004. The U.S.-Canada softwood lumber trade dispute: what we know and what we have to know. Forest Policy and Economics 6: 129-143.

Yin, R. and J. Xu. 2003. Identifying market relationships with co-integration and causality tests. Forest Policy and Economics 5: 305-315.

Zhang, D. 2000. Welfare impacts of the 1996 U.S.-Canada softwood lumber trade agreement. Canadian Journal of Forest Research 31: 1958-1967.

Zellner, A and F. Palm. 1974. Time series analysis and simultaneous equation econometric models. Journal of Econometrics 2: 17-54.

APPENDIX A: Tables

Table A-1. Results of lag order selection criteria (dataset II)

Lag	LR	FPE	AIC	SC	HQ
0	NA	7.71E-21	-20.77	-20.42	-20.64
1	509.23	3.60E-25	-30.80	-27.29	-29.47
2	146.96	9.47E-26	-32.49	-25.83	-29.97
3	169.92*	1.57E-30*	-36.72*	-25.71	-32.50
4	128.52	1.35E-26	-35.54	-34.63*	-31.83*

Note:

1. * indicates lag order selected by criterion at 5% level.
2. LR, FPE, AIC, SC, and HQ represent log likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion, respectively.
3. Given the sample size ($n = 52$), the software (EViews) only allowed up to 4 lags for the tests.

Table A-2. Results of residual tests (dataset II)

Equation	Serial Correlation $F_{AR}(3,26)$	Heteroskedasticity $F_{ARCH}(4,21)$	Normality $\chi^2(2)$
ΔEXP	0.44 [0.65]	0.06 [0.99]	1.92 [0.38]
ΔSUP	0.37 [0.29]	0.77 [0.56]	0.01 [0.99]
ΔP	0.29 [0.84]	0.27 [0.89]	2.21 [0.33]
ΔHS	0.08 [0.97]	0.32 [0.86]	1.99 [0.37]
ΔDI	1.36 [0.28]	0.25 [0.91]	4.39 [0.11]
ΔMR	0.79 [0.51]	2.06 [0.12]	1.96 [0.37]
ΔLP	2.19 [0.10]	0.33 [0.85]	2.61 [0.27]
ΔEX	2.82 [0.08]	0.62 [0.65]	2.62 [0.27]
ΔW	1.34 [0.29]	0.65 [0.64]	0.17 [0.92]
System	0.13 [0.99]	-	17.90 [0.46]

Note:

1. Values in brackets are p-value; * indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LS test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

Table A-3. Results of weak exogeneity tests (dataset II)

Equations	Weak exogeneity $H_0 : \alpha_i = 0$
EXP	11.82 [0.00]**
SUP	14.62 [0.00]**
P	20.65 [0.00]**
HS	5.51[0.14]
DI	2.89 [0.41]
MR	7.39 [0.06]
LP	1.97 [0.58]
EX	5.25 [0.15]
W	7.14 [0.07]

Note:

1. * denotes rejection of the hypothesis at the 5% significant level.
2. Weak exogeneity and variable exclusion under $r = 2$ follow χ^2 (2).

Table A-4. Results of lag order selection criteria for endogenous variables (dataset II)

Lag	LR	FPE	AIC	SC	HQ
0	NA	2.02E-06	-4.60	-4.25	-4.47
1	119.89	1.70E-07	-7.08	-6.38	-6.82
2	36.10*	1.35E-07	-7.95*	-6.54*	-6.92
3	23.90	7.34E-08*	-7.93	-6.27	-7.42*
4	11.83	7.71E-08	-7.32	-6.17	-7.27

Note:

1. * indicates lag order selected by criterion at 5% level.
2. LR, FPE, AIC, SC, and HQ represent log likelihood ratio, final prediction error, Akaike information criterion, Schwarz information criterion, and Hannan-Quinn information criterion, respectively.

Table A-5. Results of residual tests for endogenous variables (dataset II)

Equation	Serial Correlation F_{AR} (2,36)	Heteroskedasticity F_{ARCH} (4,30)	Normality χ^2 (2)
ΔEXP	1.35 [0.27]	0.18 [0.95]	2.26 [0.32]
ΔSUP	0.36 [0.71]	0.73 [0.58]	0.26 [0.88]
ΔP	1.29 [0.29]	0.74 [0.58]	2.84 [0.24]
System	0.70 [0.81]	-	11.85 [0.07]

Note:

1. Values in brackets are p-value; * indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LS test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

Table A-6. Results of residual tests for basic VEC model (dataset I)

Equation	Serial Correlation $F_{AR}(7,47)$	Heteroskedasticity $F_{ARCH}(4,46)$	Normality $\chi^2(2)$
$\Delta D(EXP)$	1.77 [0.17]	0.44 [0.78]	25.54 [0.00]*
$\Delta D(SUP)$	2.00 [0.07]	0.32 [0.87]	4.85 [0.09]
$\Delta D(P)$	1.20 [0.32]	0.12 [0.97]	1.26 [0.53]
System	1.12 [0.30]	-	4.39 [0.35]

Note:

1. Values in brackets are *p*-value;* indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LS test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

Table A-7. Results of F-tests for the basic VEC model (dataset I)

Variable		F-statistic	p-value
Endogenous	EXP_{t-1}	8.80	0.00**
	EXP_{t-2}	4.97	0.00**
	EXP_{t-3}	3.95	0.01**
	EXP_{t-4}	2.13	0.11
	P_{t-1}	1.36	0.27
	P_{t-2}	3.03	0.04**
	P_{t-3}	3.57	0.02**
	P_{t-4}	1.57	0.21
	SUP_{t-1}	4.71	0.00**
	SUP_{t-2}	2.20	0.09*
	SUP_{t-3}	1.22	0.31
	SUP_{t-4}	0.35	0.79
Exogenous	HS	6.42	0.00**
	MR	1.58	0.21
	DI	1.61	0.65
	W	1.19	0.32
	CW	0.48	0.70
	CLP	0.94	0.43
	EX	2.90	0.05**
Overall F-test: F (21, 161)		1.07	0.37

Note:

1. ** and * indicate significance at the 5% and 10% level.
2. The null hypothesis of the overall F-test is the omitted explanatory variables are zero.

Table A-8. Results of the parsimonious VEC model (dataset I)

Variable		EXP		SUP		P	
		Coefficient	*t*-value	Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.68	-5.15*	-0.18	-2.02*	-0.31	-1.31
	EXP_{t-2}	-0.39	-2.93*	-0.14	-2.46*	-0.11	-0.78
	EXP_{t-3}	-0.29	-2.60*	0.12	0.20	-0.13	-1.15
	EXP_{t-4}						
	P_{t-1}						
	P_{t-2}	-0.12	-1.17	0.01	0.19	-0.23	-2.23*
	P_{t-3}	0.01	0.03	0.16	2.46*	0.29	2.79*
	P_{t-4}						
	SUP_{t-1}	-0.20	-2.98*	-0.34	-2.32*	-0.60	-2.94*
	SUP_{t-2}	-0.11	-0.95	-0.33	-2.24*	-0.08	-1.28
	SUP_{t-3}						
	SUP_{t-4}						
Exogenous	HS	0.17	2.44*	0.26	2.83*	0.09	1.17
	MR						
	DI						
	W						
	LP						
	CLP						
	EX	0.30	1.97*	-0.17	-1.38	-0.31	-1.58
	MOU (1986:4-1991:3)	-0.09	-3.24*	0.05	2.73*	0.01	0.31
	SLA (1996:2-2001:1)	-0.04	-1.23	0.01	0.22	-0.01	-0.04
	DUM_1 (1991:1-1993:4)	0.02	0.05	-0.03	-2.20*	0.11	2.53*
	DUM_2 (1997:3-1998:4)	0.02	0.58	0.02	0.58	-0.01	-0.26
	SEA_1 (2nd quarter)	0.05	1.60	0.12	5.73*	0.13	3.79*
	SEA_2 (3rd quarter)	0.09	1.78	0.13	4.24*	0.03	0.51
	SEA_3 (4th quarter)	0.05	1.45	0.05	2.11*	-0.01	-0.31
Error-correction term		-0.55	-2.45*	-0.24	-3.54*	-0.41	-1.58

Note:

1. * indicates significance at the 5% level.

Table A-9. Results of residual tests for basic VEC model (dataset II)

Equation	Serial Correlation $F_{AR}(2,28)$	Heteroskedasticity $F_{ARCH}(4,22)$	Normality $\chi^2(2)$
$\Delta D(EXP)$	0.86 [0.43]	0.32 [0.86]	5.87 [0.06]
$\Delta D(SUP)$	0.08 [0.92]	1.12 [0.37]	0.75 [0.69]
$\Delta D(P)$	0.31 [0.74]	0.18 [0.94]	7.59 [0.02]*
System	1.52 [0.11]	-	11.47 [0.08]

Note:

1. Values in brackets are *p*-value; * indicates that the null hypothesis is rejected at 5% level.
2. Serial correlation 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 independent variables; Heteroskedasticity was tested using the F-form of the LM test; Normality of the residuals was tested with the Doornik-Hansen test (Doornik and Hansen 1994).

Table A-10. Results of F-tests for the basic VEC model (dataset II)

Variable		F-statistic	p-value
Endogenous	EXP_{t-1}	10.71	0.00*
	EXP_{t-2}	5.45	0.00*
	P_{t-1}	0.71	0.56
	P_{t-2}	3.24	0.04*
	SUP_{t-1}	4.32	0.01*
	SUP_{t-2}	1.67	0.20
Exogenous	HS	3.14	0.04*
	MR	2.30	0.11
	DI	1.73	0.18
	W	2.31	0.11
	LP	3.52	0.03*
	EX	4.80	0.00*
Overall F-test: F (6, 56)		0.96	0.46

Note:

1. * indicates significance at the 5% level.
2. The null hypothesis of the overall F-test is the omitted variables are zero coefficient.

Table A-11. Results of the parsimonious VEC model (dataset II)

Variable		EXP		SUP		P	
		Coefficient	*t*-value	Coefficient	*t*-value	Coefficient	*t*-value
Endogenous	EXP_{t-1}	-0.85	-4.71*	-0.39	-3.10*	-1.02	-2.94*
	EXP_{t-2}	-0.29	-1.63	-0.37	-2.98*	-0.18	-0.94
	P_{t-1}						
	P_{t-2}	-0.11	-0.69	0.03	1.60	-0.73	-2.14*
	SUP_{t-1}	0.39	2.00*	0.15	0.86	0.52	1.23
	SUP_{t-2}						
Exogenous	HS	0.16	2.12*	0.25	2.94*	0.20	2.68*
	MR						
	DI						
	W						
	LP	0.11	2.98*	-0.01	-1.13	-0.02	-1.19
	EX	0.29	1.98*	-0.41	-3.54*	-0.99	-3.17*
	MOU (1989:1-1991:3)	-0.01	-0.12	0.05	0.20	-0.02	-0.88
	SLA (1996:2-2001:1)	-0.05	-1.05	-0.01	-1.23	-0.03	-1.01
	SEA_1 (2nd quarter)	0.01	2.04*	0.06	2.60*	-0.05	-1.48
	SEA_2 (3rd quarter)	0.07	1.44	-0.02	-1.59	-0.09	-2.08*
	SEA_3 (4th quarter)	0.09	2.48*	-0.01	-1.40	-0.09	-1.25
Error-correction term		-0.50	-2.47*	-0.43	-3.41*	-0.46	-2.42*
Multivariate tests		$F_{ar}(18,76) = 0.97$; $F_{het}(60,104) = 0.53$; $\chi^2(6) = 11.46$					

Note:

1. * indicates significance at the 5% level.
2. Multivariate tests for serial correlation, heteroskedasticity, and normality were not rejected at the 5% significant level.

Table A-12. Results of market and welfare impacts under MOU (case I)

		1986:4-1987:3			1986:4-1991:3		
		With MOU	Without MOU	Policy impact	Total change	Annual change	% change
Market impact	Consumption (MMBF)	49,237.0	49,697.3	-460.3 (-0.9)	-2,211.1	-442.2	-0.9
	Production (MMBF)	35,668.6	35,043.4	+625.2 (+1.8)	+2,996.1	+599.2	+1.8
	Canadian exports (MMBF)	13,568.4	14,653.9	-1,085.5 (-8.0)	-5,207.2	-1,041.4	-8.0
	Price (U.S. $/MBF)	205.3	181.3	+24.2 (+11.7)	-	+21.9	+11.8
Welfare impact	U.S. consumer surplus	63,176.5	64,363.2	-1,186.7 (-1.8)	-5,191.8	-1,038.4	-1.9
	U.S. producer surplus	6,773.5	5,925.3	+848.2 (+14.3)	+3,700.1	+740.0	+14.4
	U.S. net impact	69,950.0	70,288.6	-338.5 (-0.5)	-1,491.8	-298.4	-0.5
	CA producer surplus	2,367.6	2,522.9	-155.2 (-6.2)	-697.2	-139.4	-6.4
	Export tax revenue	417.8	0.0	+417.8	+1,825.0	+365.0	-
	CA net impact	2,785.5	2,522.9	+262.6 (+10.4)	+1,127.9	+225.6	+10.2
	Deadweight loss	-	-	-75.9	-363.9	-72.8	-

Note:

1. All values for welfare impact are in million 1982 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset II are adopted (demand: -0.08, domestic supply: 0.15, and Canadian exports: 0.15).
3. Deadweight loss is the sum of U.S. and Canadian net impacts.

Table A-13. Results of market and welfare impacts under MOU (case II)

		1986:4-1987:3			1986:4-1991:3		
		With MOU	Without MOU	Policy impact	Total change	Annual change	% change
Market impact	Consumption (MMBF)	49,237.0	49,388.1	-151.1 (-0.3)	-753.8	-150.8	-0.3
	Production (MMBF)	35,668.6	34,598.5	+1,070.1 (+3.0)	+5,104.2	+1,020.8	+3.0
	Canadian exports (MMBF)	13,568.4	14,789.6	-1,221.2 (-9.0)	-5,858.1	-1,171.6	-9.0
	Price (U.S. $/MBF)	205.3	205.3	0.0	-	0.0	0.0
Welfare impact	U.S. consumer surplus	63,176.5	63,372.4	-193.9 (-0.3)	-888.3	-177.7	-0.3
	U.S. producer surplus	6,773.5	6,570.3	+203.2 (+3.1)	+881.4	+176.3	+3.1
	U.S. net impact	69,950.0	69,940.7	+9.3 (+0.1)	+6.9	+1.4	+0.1
	CA producer surplus	2,367.6	2,883.1	-417.8 (-17.9)	-2,288.4	-457.7	-18.1
	Export tax revenue	417.8	0.0	+417.8	+1,825.0	+365.0	-
	CA net impact	2,785.5	2,883.1	+97.7 (-3.4)	-463.4	+92.7	-3.7
	Deadweight loss	-	-	-88.3	-470.3	-94.1	-

Note:

1. All values for welfare impact are in million 1982 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset II are adopted (demand: -0.08, domestic supply: 0.15, and Canadian exports: 0.15).
3. Deadweight loss is the sum of U.S. and Canadian net impacts.

Table A-14. Results of market and welfare impacts under SLA (dataset II)

		1996:2-1997:1			1996:2-2001.1		
		With SLA	Without SLA	Policy impact	Total change	Annual change	% change
Market impact	Consumption (MMBF)	49,651.0	49,934.9	-289.9 (-0.6)	-1,347.1	-268.4	-0.5
	Production (MMBF)	32,165.2	31,820.3	+344.9 (+1.1)	+1,755.8	+351.2	+1.0
	Canadian exports (MMBF)	15,719.0	16,347.8	-628.8 (-4.0)	+3,103.0	-620.6	-4.0
	Other exports (MMBF)	1,766.8	1,577.4	+189.4 (+10.7)	+1,149.3	+229.9	+9.7
	Price (U.S. $/MBF)	431.7	400.9	+30.9 (+7.2)	-	+23.8	+6.5
Welfare impact	U.S. consumer surplus	133,968.2	135,504.7	-1,536.4 (-1.1)	-6,221.8	-1,244.4	-1.0
	U.S. producer surplus	12,844.6	11,857.4	+987.2 (+8.3)	+4,277.9	+855.6	+7.5
	U.S. net impact	146,812.9	147,362.1	-549.3 (-0.4)	-1,943.8	-388.8	-0.3
	CA producer surplus	3,274.5	2,877.2	+397.3 (+13.8)	+1,504.4	+300.9	+12.4
	Other producer surplus	381.4	316.2	+65.2 (+20.6)	+333.3	+66.7	+4.1
	CA net impact	3,655.8	3,193.4	+462.5 (+14.5)	+1,837.7	+367.5	+13.2
	Deadweight loss	-	-	-86.8	-106.2	-21.2	-

Note:

1. All values for welfare impact are in million 1997 U.S. $; Values in parentheses are percentages.
2. Price elasticities estimated from dataset II are adopted (demand: -0.08, domestic supply: 0.15, and Canadian exports: 0.15).
3. Deadweight loss is the sum of U.S. and Canadian net impacts.
4. Other exports (producer surplus) are the volume (gain) of Canadian exporters in non-quota constrained provinces, such as New Brunswick and Nova Scotia.

APPENDIX B: Figures

Figure B-1. Co-integration vectors and recursive estimates of the eigenvalues (dataset II)

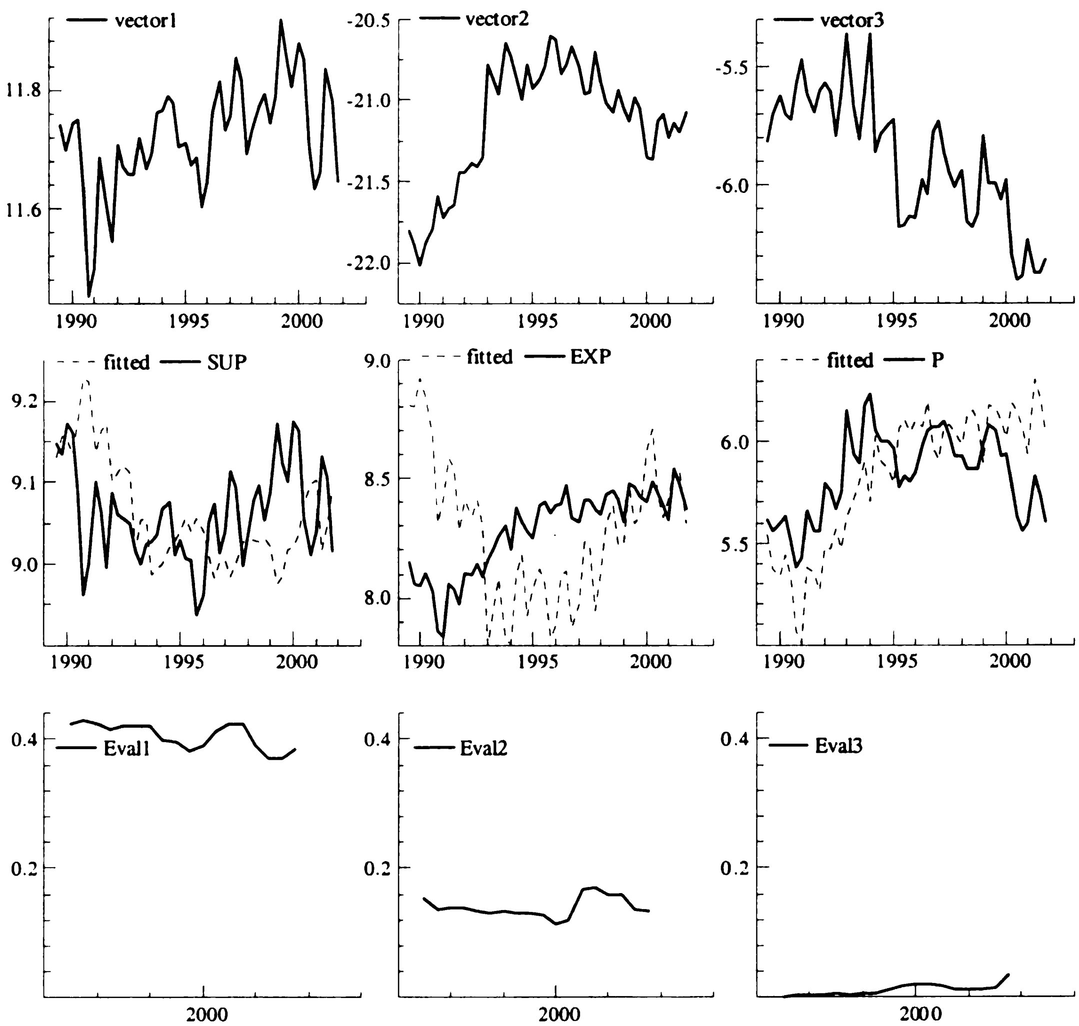


Figure B-2. Fitted and actual values and diagnostics for basic VEC model (dataset I)

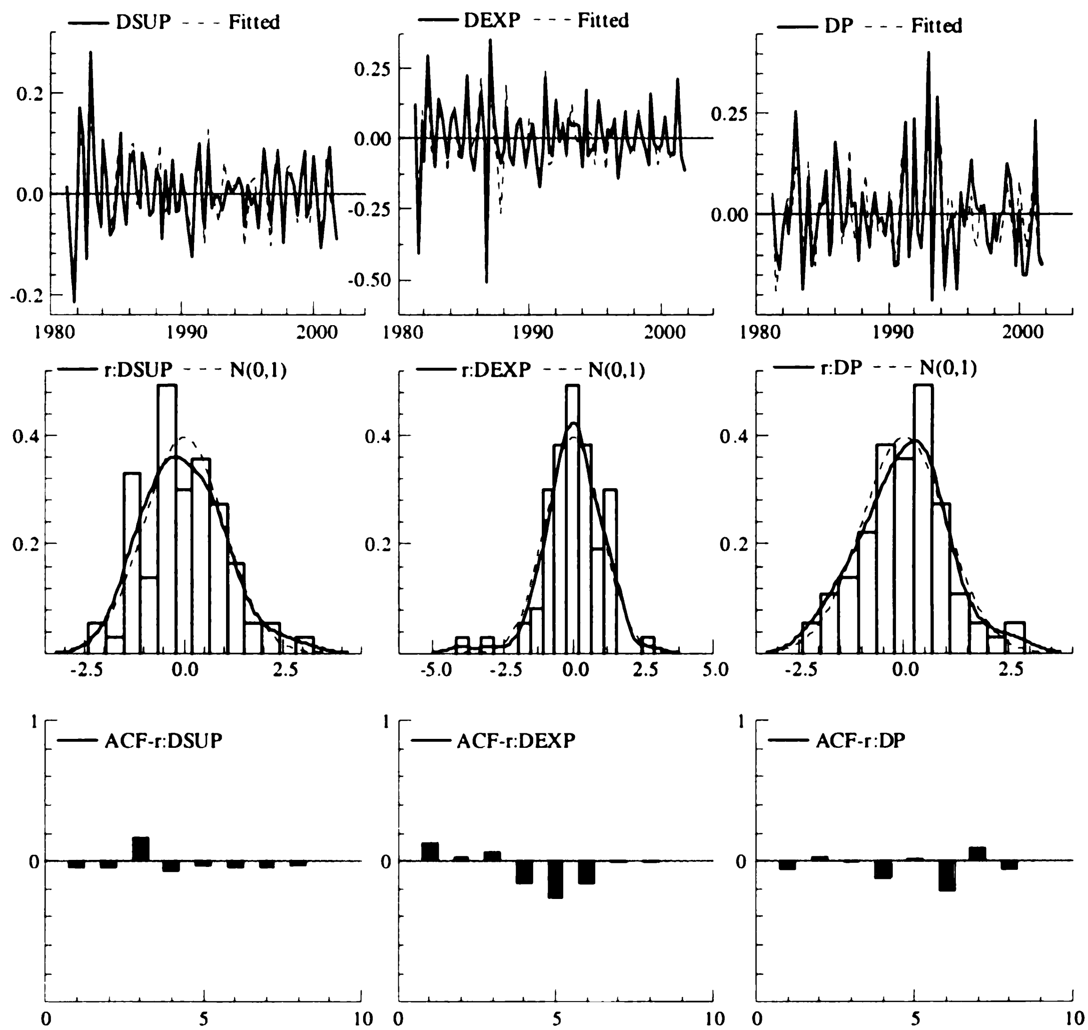


Figure B-3. One-step residuals and Chow test of basic VEC model (dataset I)

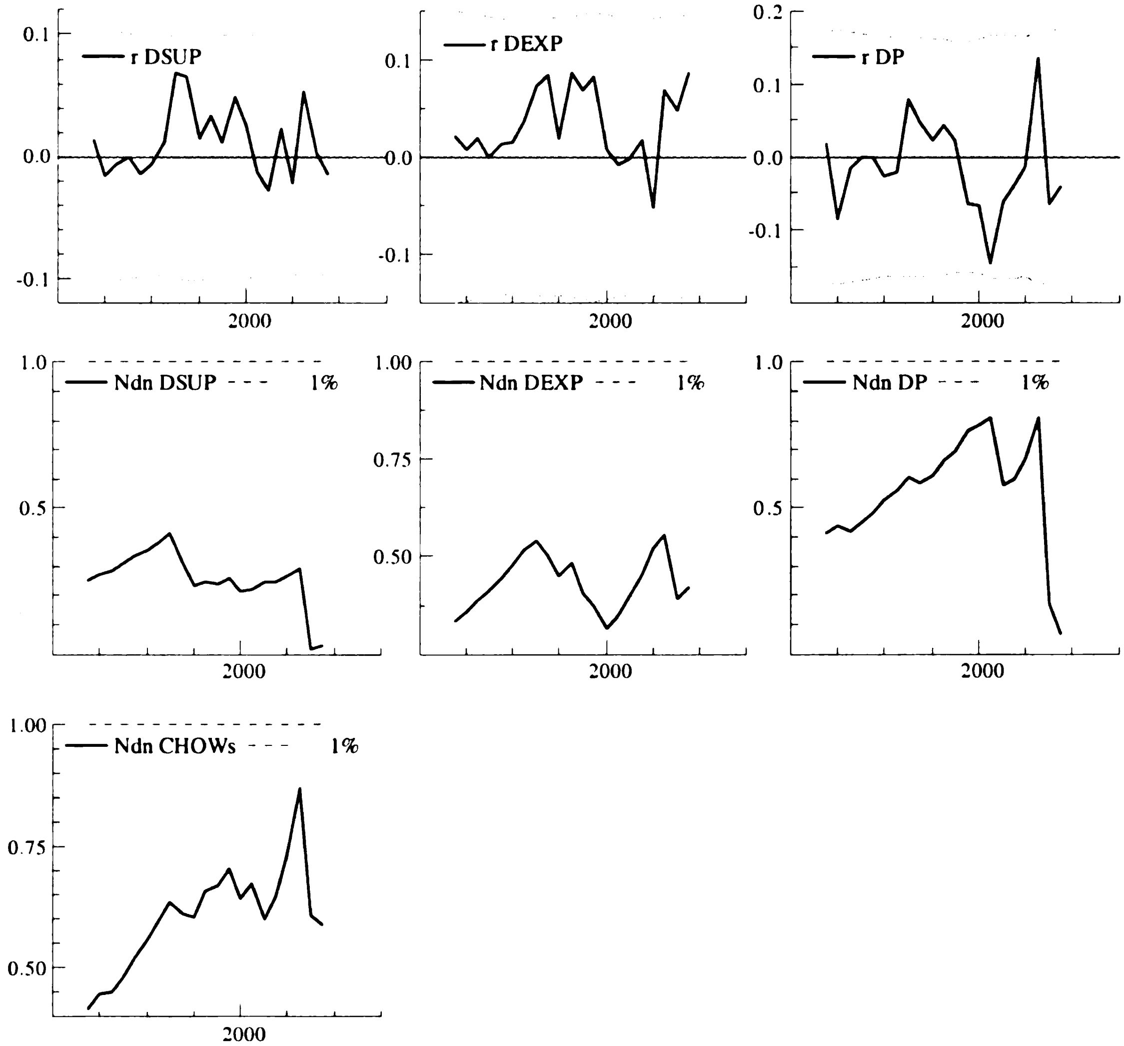


Figure B-4. Fitted and actual values and diagnostic for SPVEC model (dataset I)

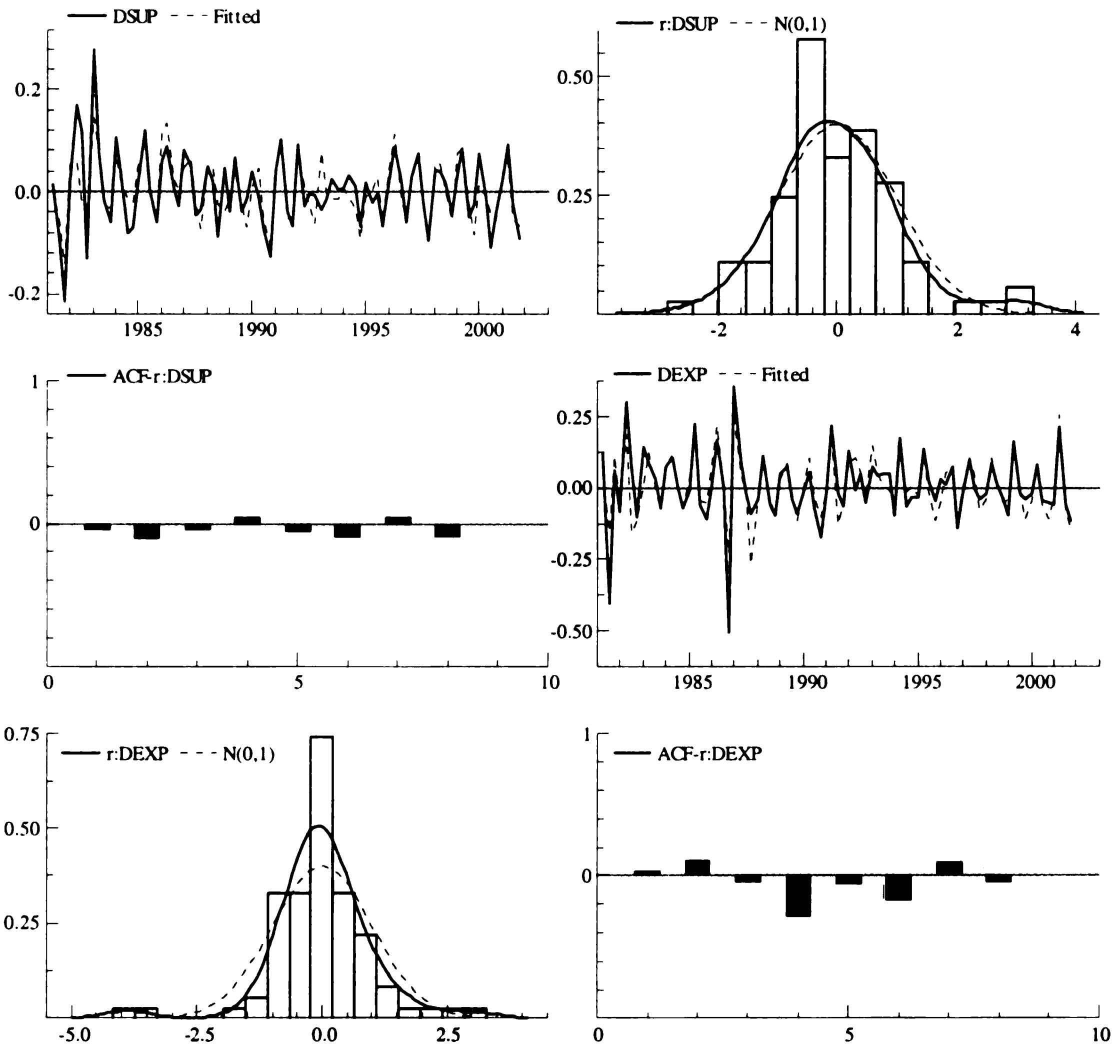


Figure B-5. One-step residuals and Chow tests of SPVEC model (dataset I)

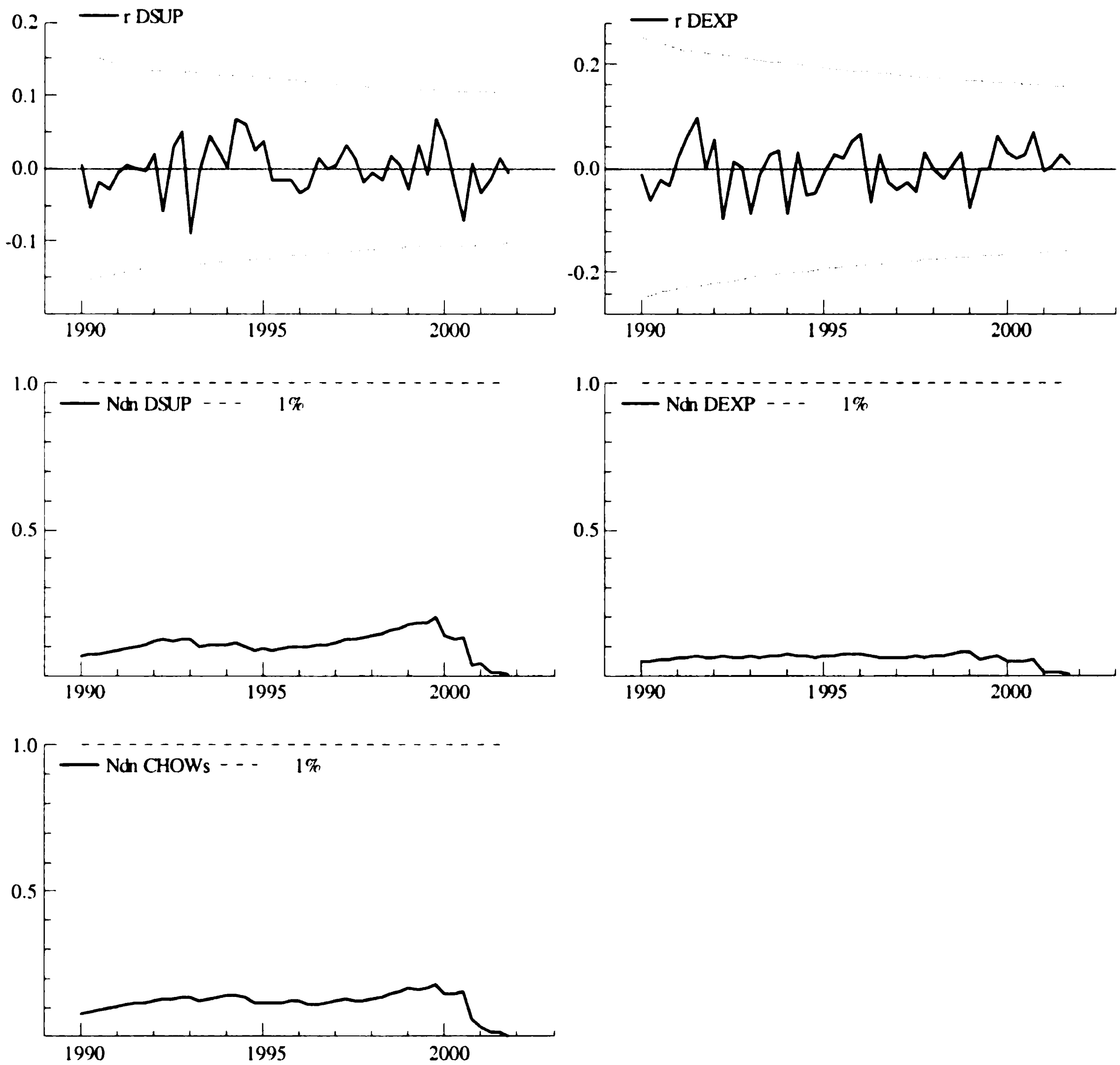


Figure B-6. Fitted and actual values and diagnostics for basic VEC model (dataset II)

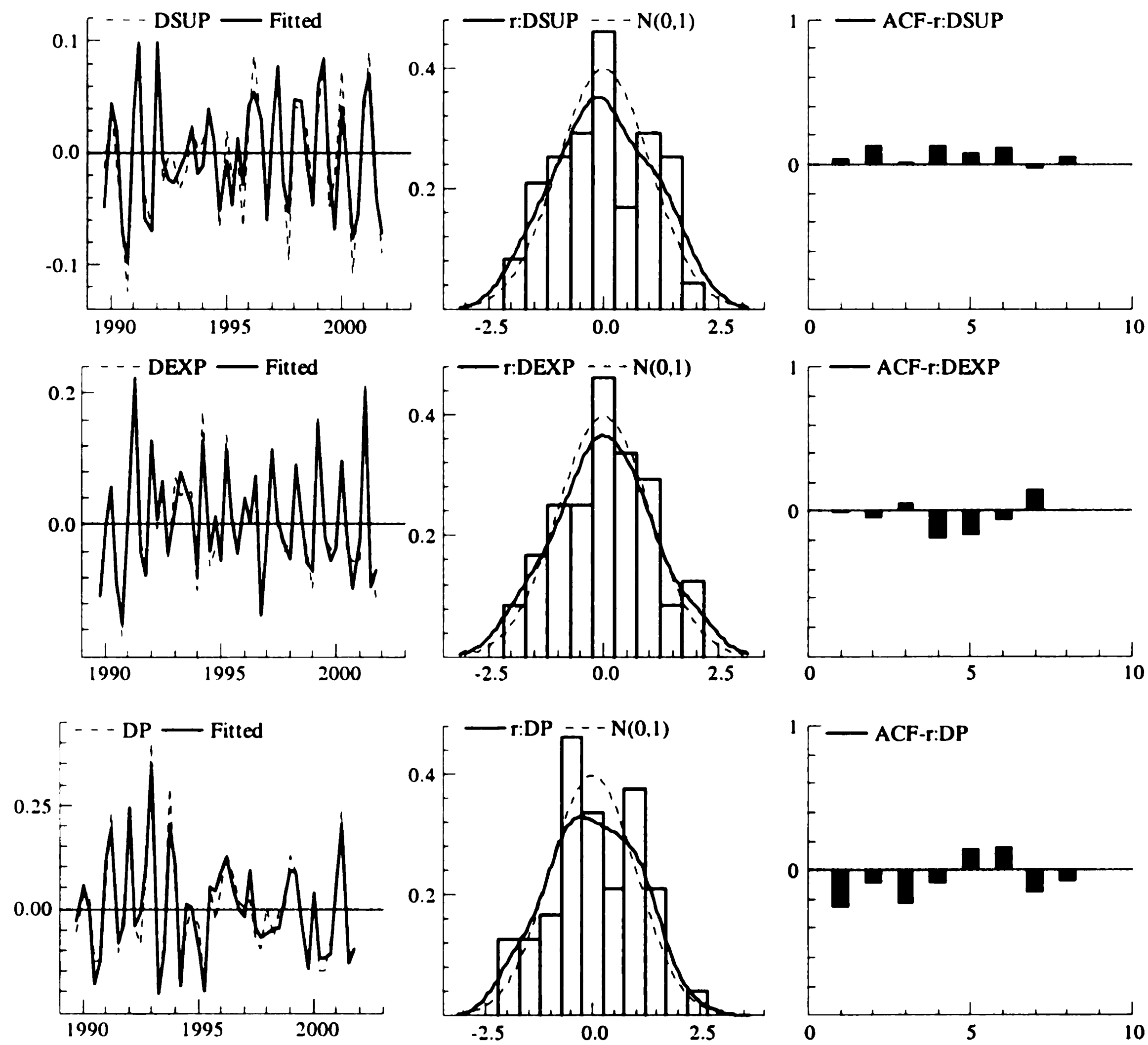


Figure B-7. One-step residuals and Chow tests for basic VEC model (dataset II)

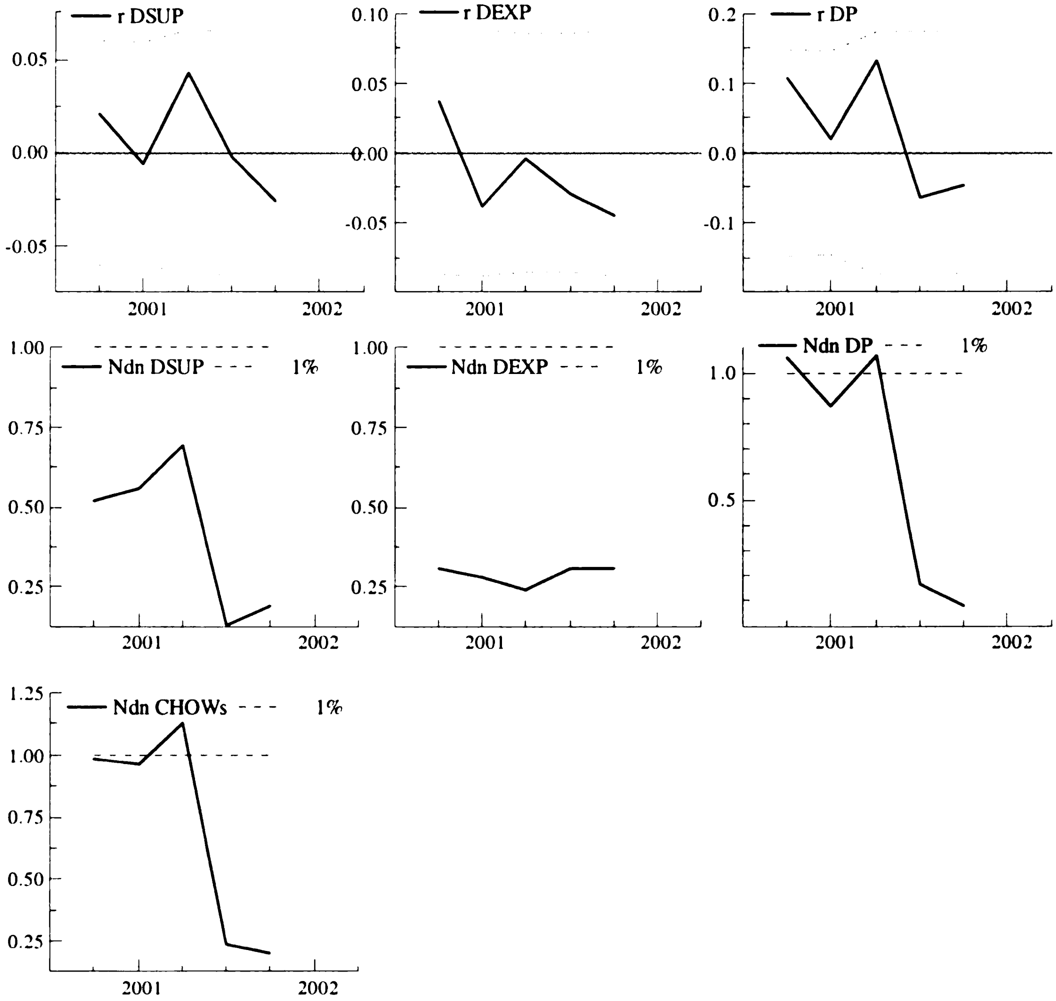


Figure B-8. Fitted and actual values and diagnostics for SPVEC model (dataset II)

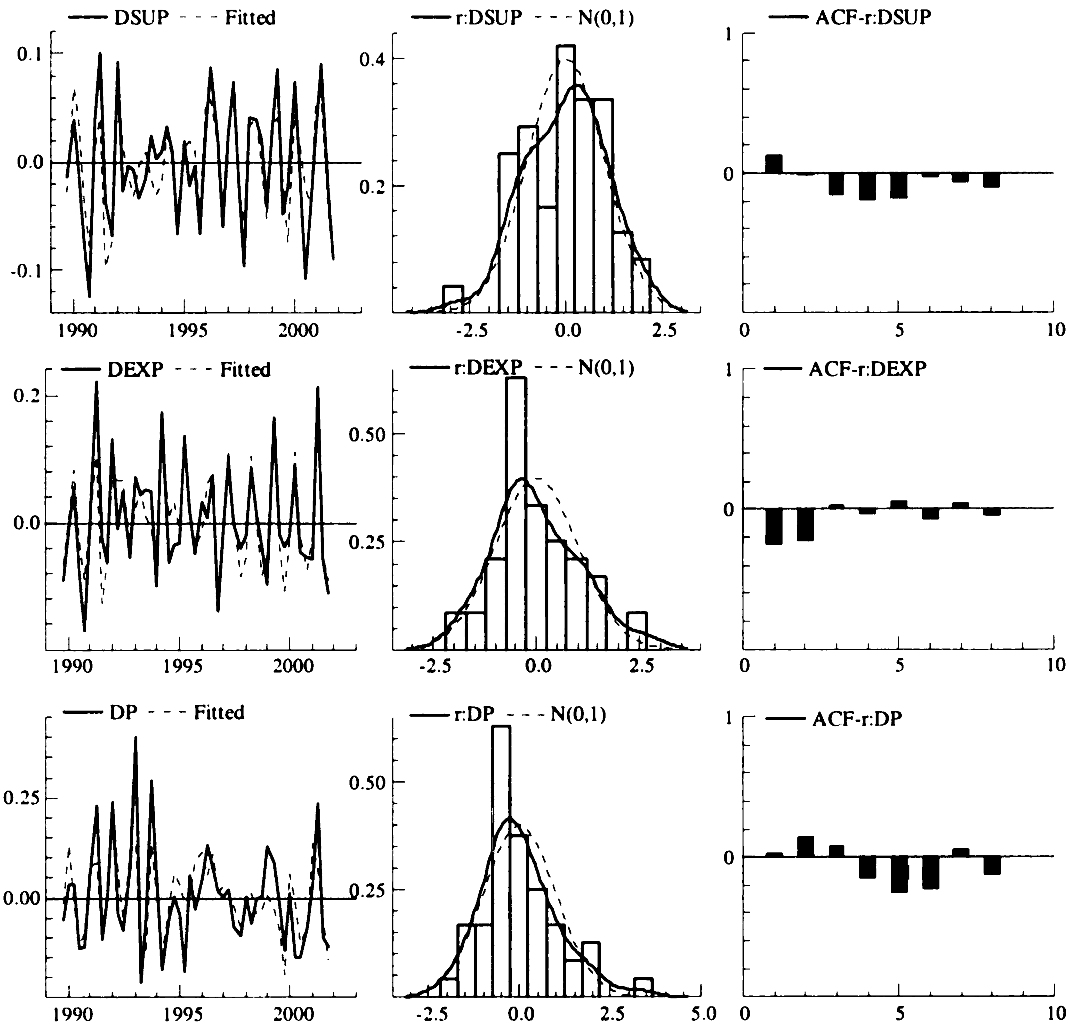


Figure B-9. One-step residuals and Chow tests for SPVEC model (dataset II)