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# Identifying the inter-market relationships of forest products in the Pacific Northwest with cointegration and causality tests

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**Abstract**

We identify various inter-market relationships of forest products using cointegration and causality tests together. Of the six Douglas fir domestic sawlog, export sawlog, and lumber markets in the Pacific Northwest, we find that the two log markets and the two lumber markets are integrated, respectively. However, the two export log markets are not, nor is any cross-grade combination. In conjunction with cointegration restrictions, our causality tests demonstrate that export and lumber prices lead the movement of domestic sawlog prices; and similarly, the movements of domestic lumber prices follow the movements of export log prices. A close examination further reveals that export log prices for Region 1 lead the price formation process in all the lumber markets and log markets. We believe that these results have significant implications for understanding and thus dealing with forest products market behavior and price forecasting.



*Keywords:* Inter-market relationship; Cointegration; Causality; Weak exogeneity; Price forecasting

## 1. Introduction

Market relationships encompass intra- and inter-market types, both primarily reflected in the underlying price movements (Hamilton, 1994; Harvey, 1993; Pankranz, 1990). It has become standard practice in economics to analyze these market relationships using price series. To characterize intra-market relationships, analysts use univariate autoregressive moving-average, or ARMA, models based on a single price series (Box and Jenkins, 1976; Hamilton, 1994; Pankratz, 1991), while for inter-market relationships analysts use vector autoregression, or VAR, models based on multiple price series (Granger, 1969; Engle and Granger, 1987; Johansen, 1995). Even though these market relationships are paramount to effective market modeling and price forecasting in the forest sector, limited efforts have been made to examine them (Yin, 2001; Prestemon and Holmes, 2000; Hänninen, 1998; Buongiorno and Uusivuori, 1992;

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Buongiorno et al., 1988). The goal of this paper is thus to identify various inter-market relationships for forest products.

Inter-market relationships comprise spatial relationships between different market segments of the same product, grade relationships between different products in the same region, and vertical relationships between products of different manufacturing stages. In the context of a forest economy, an example of a spatial relationship is that of sawlog prices between two regions; a grade relationship is that between sawlog and pulpwood prices, or domestic and export sawlog prices in the same region; and a vertical relationship is that of sawlog and lumber prices again in the same region. As shown later, these spatial, grade, and vertical relationships can be examined in terms of their degrees of market integration and precedence (Hamilton, 1994; Pankratz, 1991). Thanks to the pioneering work of Granger (1969, 1988) and others, it is now accepted that market integration can be analyzed with the cointegration test, and market precedence with the causality test.

Intuitively, while individual price series may behave like random walks, over the long run they tend to drift in similar fashion. If this is indeed the case, then we can state that the underlying price series are cointegrated; hence, the corresponding markets are integrated (Hamilton, 1994). Even without cointegration, however, an association may exist between two non-stationary series. That is, one series may have the power to predict or even determine another. In this situation, we say that one price series 'Granger-causes' another series; thus, there is a market precedence between them (Hamilton, 1994; Granger, 1988).[2]

From the above discussion, it appears that cointegration implies causality. More specifically, if two non-stationary series are cointegrated, then a linear combination of them should reduce to a stationary and presumably noise process (Harvey, 1993; Granger, 1988). Therefore, one may be able to predict the other. On the other hand, even if one series is able to predict the other, cointegration is not given. Also, cointegration characterizes the long-run relationship between the underlying series, whereas causality features the short-run relationship (Granger, 1988). Therefore, the former has stricter requirements than the latter for the series of concern, and the latter occurs more frequently than the former. Since the distinctions between cointegration and causality are more complex, however, caution should be taken when making assumptions.

The objective of this paper is to identify various relationships of forest products markets by examining the degree of their integration and precedence. To that end, we will conduct cointegration and causality tests across spatial, grade, and vertical dimensions using price series from the Pacific Northwest (PNW). As the PNW is a major forest area in the US and its forest sector has a significant impact on the regional economy, it is a natural locale for understanding market relationships. Secondly, the tremendous market shifts in the region since the late 1980s, induced by the spotted owl debate and the Asian financial crisis (Yaffe, 1994; Sohngen and Haynes, 1994; Murray and Wear, 1998), offer greater interest in examining relationships between the price series.

Cointegration and causality tests have appeared in the forest economics literature. Using the bivariate Dickey–Fuller procedure, Buongiorno and Uusivuori (1992) found that the law of one price in US pulp and paper export markets could not be rejected in most pairs of price series analyzed. Adopting the multivariate Johansen procedure to evaluate long-run equilibrium relationships among four US regional lumber markets, Jung and Doroodian (1994) concluded that the law of one price prevailed. Based on both cointegration and causality tests, Murray and Wear (1998) detected a structural break in the relationship between lumber price series in the PNW and the South around the time when federal harvest restrictions were imposed, resulting in a more integrated market.

The Granger-causality test used by Buongiorno et al. (1985) did not detect any significant influence of National Forest timber sales on the price of lumber. In another application of causality test, Uri and Boyd (1990) suggested that the demand

[2] Notably, the reason for us to focus on non-stationary series is that most economic series are indeed non-stationary; otherwise, we can simply use standard econometric techniques to analyze them (Hamilton, 1994).

for softwood lumber is indeed strongly connected to prices, and there exists a national market for it. More recently, an interesting book, *Modern Time Series Analysis in Forest Products Markets* (Abildtrup et al., 1999), has documented the extensive applications of cointegration, causality, and other related analyses to forest products markets in Europe and the United States.

While both cointegration and causality tests have been used to examine market relationships of forest products, they have rarely been used in combination. Further, previous studies tend to deal with single rather than multiple dimensional market relationships, and they did not give sufficient attention to the specification of a VAR used for the tests. Building on these studies, this work attempts to disentangle multiple market relationships at different levels. It is hoped that this analysis will shed new light on market relationships for forest products, which should be beneficial to market modelers, business analysts, and wood products manufacturers and consumers.

## 2. Procedures

The mathematical form of a VAR is

$$y_t = A_1 y_{t-1} + \cdots + A_p y_{t-p} + B x_t + \varepsilon_t \tag{1}$$

where $y_t$ is an $n$-vector of non-stationary $I(1)$ variables, $x_t$ is a $d$-vector of deterministic variables, $A_1,\ldots, A_p$ and $B$ are matrices of coefficients to be estimated, and $\varepsilon_t$ is a vector of innovations that may be contemporaneously correlated with each other but are uncorrelated with their own lagged values and other right-hand side variables. In our case, $y_t$ consists of price series, and $x_t$ may include variables like intercept and time trend.

### 2.1. Cointegration

An $(n\times 1)$ vector of price series, $y_t$ is cointegrated if each of the elements of $y_t$ must be $I(1)$ individually—non-stationary with a unit root, while some linear combination of the series $a'y_t$ is stationary, or $I(0)$ for some nonzero $(n\times 1)$ vector, $a$. Cointegration means that even though many developments can cause permanent changes in the individual elements of $y_t$, a long-run equilibrium relationship ties these individual series together, represented by the linear combination $a'y_t$. A test of the hypothesis that $z_t = a'y_t$ is $I(1)$ is thus equivalent to a test of the hypothesis that $y_t$ is not cointegrated. If the null hypothesis that $z_t$ is $I(1)$ is rejected, we will conclude that $z_t$ is stationary, or that $y_t$ is cointegrated.

In this paper, we utilize the approach of Johansen (1995) in testing the null hypothesis that $z_t$ is $I(1)$, which is capable of dealing with simultaneity bias. The test statistic is obtained via a maximum likelihood estimation method. For the Johansen test, we can rewrite the above VAR as:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + B x_t + \varepsilon_t \tag{2}$$

with

$$\Pi = \textstyle\sum_{i=1}^{p} A_i - I, \quad \Gamma_i = -\sum_{j=i+1}^{p} A_j.$$

Granger's representation theorem asserts that if the coefficient matrix $\Pi$ has reduced rank $r<n$, then there exist $n\times r$ matrices $\alpha$ and $\beta$ each with rank $r$ such that $\Pi = \alpha\beta'$ and $\beta' y_t$ is stationary (Engle and Granger, 1987). Here, $r$ is the number of cointegrating relations and each column of $\beta$ is a cointegrating vector. For $n$ endogenous non-stationary variables, there can be from 0 to $n-1$ linearly independent, cointegrating relations.

A critical consideration in a cointegration test is the structure of the model to be specified. Previous analyses of forest products markets have largely ignored the potential impact of alternative specifications, particularly the presence of an intercept and/or a deterministic trend in a cointegrating equation. However, the presence of these variables can affect the asymptotic properties of the testing statistics, making the Likelihood Ratio for the reduced rank test lack the usual $\chi^2$ distribution (Johansen, 1995). To obtain valid results, we will explore a variety of specifications of the VAR.

### 2.2. Causality

According to Granger (1969), Sims (1972), the question of whether one series, $y_1$, causes another,

$y_2$, can be answered as follows. First, we examine how much of the current $y_1$ can be explained by the past values of $y_1$. Then, we see whether adding lagged values of $y_2$ can improve the explanation. $y_1$ is said to be Granger-caused by $y_2$ if $y_2$ helps in the prediction of $y_1$, or equivalently, if the coefficients on the lagged variables of $y_2$ are statistically significant.

Using this definition, an econometric implementation of the Granger-causality test can be conducted as follows. First, we estimate

$$y_{1t} = c_1 + \alpha_1 y_{1t-1} + \alpha_2 y_{1t-2} + \cdots + \alpha_p y_{1t-p} + \beta_1 y_{2t-1} + \beta_2 y_{2t-2} + \cdots + \beta_p y_{2t-p} + u_t \quad (3)$$

by OLS. Then, we conduct an F-test of the null hypothesis $H_0\colon \beta_1 = \beta_2 = \cdots = \beta_p = 0$. One way to do this test is to calculate the sum of squared residuals from Eq. (3), $RSS_1 = \sum_{t=1}^{T} \hat{u}_t^2$, and compare that with the sum of squared residuals of a univariate autoregression for $y_t$, $RSS_0 = \sum_{t=1}^{T} \hat{e}_t^2$. If

$$S = \frac{(RRS_0 - RSS_1)/p}{RSS_1/(T - 2p - 1)} \quad (4)$$

is greater than the critical value for an $F(p, T - 2p - 1)$ distribution, then we reject the null hypothesis that $y_2$ does not Granger-cause $y_1$ (Quantitative Micro Software, 1998).

A few issues should be noted in carrying out the Granger-causality test. First, it is a bivariate procedure and thus must be used between two series. Second, the test is normally interpreted as a test of whether $y_2$ helps forecast $y_1$, rather than a test of whether $y_2$ causes $y_1$ in a more common sense of the term, although it is not impossible to go by the latter explanation if the test is formulated under the proper circumstances (Hamilton, 1994). Therefore, Granger causality measures precedence and information content but does not by itself indicate causality. Also, while a two-way causality is more frequent, it is rarely the case that the degree at which $y_2$ Ganger-causes $y_1$ is the same as that at which $y_1$ Granger-causes $y_2$. Finally, since results of causality tests can be sensitive to the choice of lag length ($p$) (Harvey, 1993), we should try multiple lags to find the most statistically sensible one.

From above definitions, it becomes clear that 'cointegration is concerned with the long-run equilibrium, whereas the causality in mean is concerned with short-run forecastability' (Granger, 1988, p. 203). That is, as long as two series track each other in the long run, they are cointegrated. It is so despite the possibility that one may drift away from the other temporarily. However, cointegration does not indicate anything about what series leads or follows the other in the short run, which is what causality means. In other words, causality implies that the movement of one series should induce or be followed by the movement of the other series. Notice that no simultaneous comovement is required here. Additionally, causality does not necessarily restrict a linear combination of the two non-stationary series to reduce to a stationary process. Thus, it is likely that cointegration of two time series will lead to their causality, but not the other way around. These differences should be recognized, and distinguishing them can add analytic intrigue and empirical appeal to research.Also, it must be made clear that, if a pair of series are cointegrated, the causality testing procedure discussed above may be no longer appropriate. The reason is that, in this case, the two series are generated by an 'error-correction' mechanism; without correcting the error using the identified cointegrating relationship, the causality testing model could be mis-specified (Granger, 1988). To rectify this problem, the causality test must be conducted in a vector error correction, or VEC, framework, in which the cointegrating restriction is imposed (Engle and Granger, 1987). In the literature, this is also called weak exogeneity test (Johansen, 1995; Doornik and Hendry, 1994).

Finally, it is worth mentioning that market relationships identified with these tests are based on statistical inferences, which, in reality, may not always be clear-cut or conform with theory. Therefore, an inference should not be drawn without thorough inquiry. In other words, we ought to exhaust all possibilities in testing before arriving at any conclusions.

## 3. Data

The specific market relationships considered in this paper are those among Douglas fir domestic sawlogs, export sawlogs, and lumber in the PNW. Log price information is from the *Log Lines 1998 Statistical Yearbook* (Arbor-Pacific Forestry Services Inc., 1998), and lumber price data are from the *Random Lengths Forest Product Price and Market Statistics Yearbook* (Random Lengths Publications, Inc., 1995–99). Because of the lack of consecutive price observations for many log grades, we had to concentrate on #2 domestic sawlogs and Japan 14 export logs; even for these series, full observations are available only for Regions 1 (covering 13 counties in Washington) and 3 (including 11 counties along the Columbia River). Fortunately, Douglas fir #2 domestic logs and Japan 14 export logs are two dominant grades, and Regions 1 and 3 are two large markets for them.[3] For this study, 108 monthly observations were included between January 1989 and December 1997.

Timber sales prices from the National Forests are more aggregated temporally (quarterly or yearly) and spatially (westside and eastside), and they do not delineate prices for different grades (Ruderman and Haynes, 1986). Also, they tend to under-represent real market prices when few bidders are involved in transactions (Sohngen and Haynes, 1994) or when below-cost sales occur (Yaffe, 1994). By contrast, price listings in Arbor-Pacific's reports represent information gathered from a minimum of three sources in a smaller geographic area for a specific grade. Hence, it appears that the Arbor-Pacific prices provide 'a reliable source of market information from private timber sales' (*Log Lines 1998 Statistical Yearbook*, p. 1).

Accordingly, our monthly lumber prices feature those for 2×4–8′ Douglas fir stud and 2×6 #2 and Better dimension lumber over the same period of time—1989:1–1997:12. They are f.o.b prices in Portland, Oregon. 2×4–8′ stud and 2×6 #2 and Better dimension are two major lumber grades. Together, they should be able to reflect price dynamics of various lumber grades. Also, as shown in the *Random Lengths 1999 Yearbook*, price movements for various lumber grades appear to be closely linked, making it redundant to use more series.

Fig. 1 shows that the two price series for 2×4–8′ stud and 2×6 #2 and Better dimension lumber track each other closely. So, we may conjecture that the two price series are cointegrated, and thus the two markets are integrated. Similarly, price series for domestic logs and export logs in the two regions appear to be cointegrated, respectively. This does not appear to hold true for any other combination of cross-grade price series (prices for domestic and export logs in Region 3, for instance). Nevertheless, similar fluctuations among all the price series suggest the possible presence of causality relationships. Our remaining task is to determine the validity of these propositions.

## 4. Results

As a precursor to the cointegration test, we first examined whether or not each of the six price series is an $I(1)$. Our Dickey–Fuller tests showed that, indeed, these price series are non-stationary with a unit root.[4] Then, we estimated the cointegrating relationships of the whole system of the six series, and we did so with specifications of various combinations of lag length, data trend, and intercept and time trend in cointegrating equations. It turned out that only two cointegrating relations were identified in most specifications. Naturally, we suspected that price pairs from two of the three market types had a higher likelihood to be cointegrated. Therefore, we moved onto the next step—searching for the specific cointegration relationships in a pair-wise manner.

Table 1 summarizes the results of these pair-wise cointegration tests. It can be seen that for the two domestic sawlog markets and the two lumber markets, estimated cointegrating relations range from zero to two. Since two price series can have

[3] For detailed definitions of domestic and export log grades, and price regions, see *Log Lines Statistical Yearbook*, published by Arbor-Pacific Forestry Services Inc.

[4] For saving space, we decided not to report the results of the unit root test.

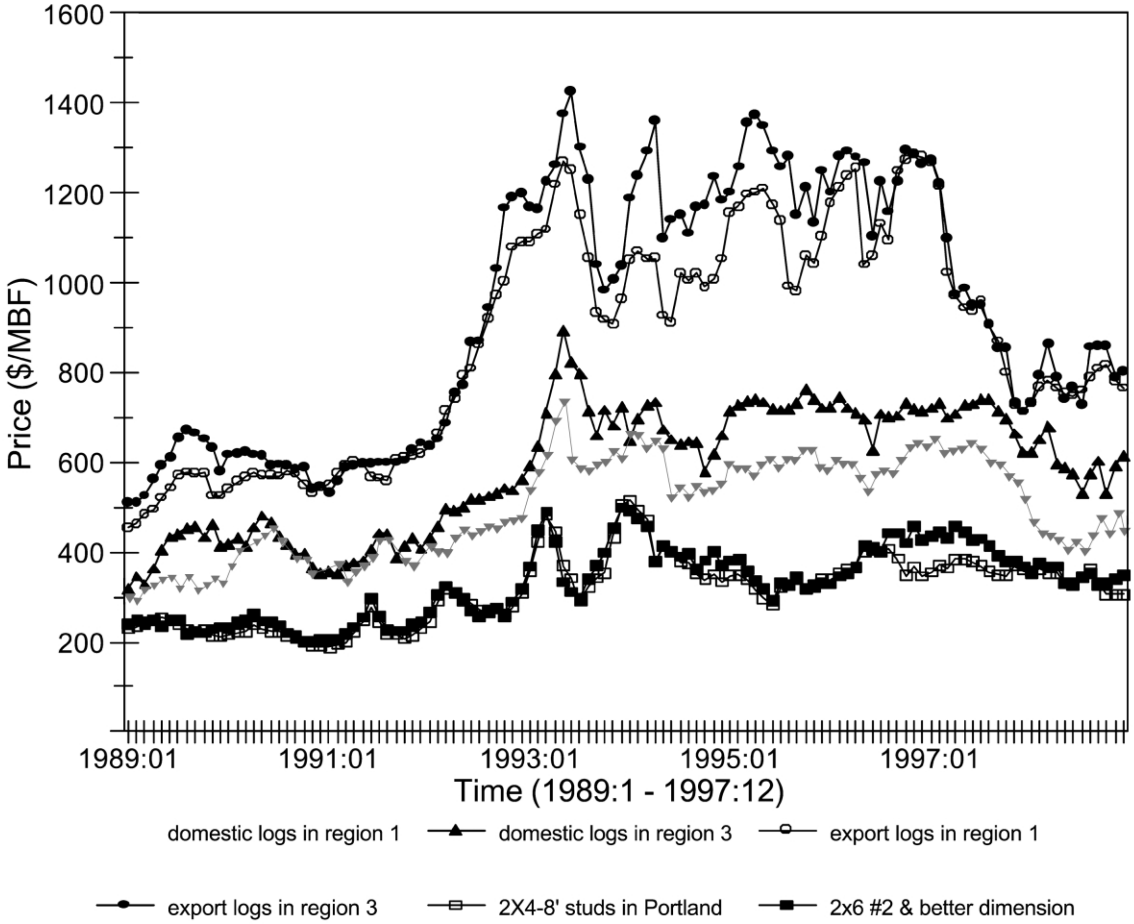


Fig. 1. Log and lumber prices in the Pacific Northwest.

only one cointegrating relation, this means that, at the least, the scenarios with two cointegrating relations are mis-specified, unless the two series are both stationary. One cointegration relationship between the two lumber price series was corroborated by 10 of the 15 specifications, indicating that it is highly likely that the two lumber grades are integrated.[5] It also appears probable that the two domestic log regions are integrated. One may argue that this inference is problematic, because results from 11 out of the 15 specifications suggest otherwise. As mentioned earlier, however, cointegration test is sensitive to how the model is specified, and a mis-specified model tends to result in the rejection of the null hypothesis. Although in only four cases where a cointegration relationship was found, we lean towards its acceptance with caution. Of course, if it is deemed that the two log regions are not integrated, then it is impossible to further perform the weak exogeneity test; the standard Granger-causality test would be sufficient between them.[6]

[5] Some acute analysts would quickly say: You do not have to report results from so many alternative specifications; simply give the one based on a selected information criterion (e.g. AIC) or the maximal value of likelihood functions. The fact is that these statistics can be very close even for distinctively different outcomes.

[6] For multiple series, however, the weak exogeneity test is feasible so long as some of them are cointegrated.

Table 1
Results of Johansen cointegration tests of forest products markets for Douglas fir in the Pacific Northwest

| Lag | Data trend<br>Intercept<br>Trend | None<br>None<br>None | None<br>Yes<br>None | Linear<br>Yes<br>None | Linear<br>Yes<br>Yes | Quadratic<br>Yes<br>Yes |
|---|---|---|---|---|---|---|
| | Lumber prices between 2×4 stud and 2×6 #2 and Better dimension | | | | | |
| 1 | | 1 | 1 | 2 | 1 | 2 |
| 2 | | 1 | 1 | 2 | 1 | 2 |
| 3 | | 1 | 1 | 1 | 1 | 2 |
| | Domestic sawlog price series between Regions 1 and 3 | | | | | |
| 1 | | 1 | 1 | 2 | 0 | 2 |
| 2 | | 0 | 0 | 1 | 0 | 0 |
| 3 | | 0 | 0 | 1 | 0 | 0 |
| | Export sawlog price series between Regions 1 and 3 | | | | | |
| 1 | | 0 | 0 | 0 | 0 | 0 |
| 2 | | 0 | 0 | 0 | 0 | 0 |
| 3 | | 0 | 0 | 0 | 0 | 0 |

(1) The alternative specifications were based on Johansen (1995, pp. 80–84) and estimated with EViews 3 (Quantitative Micro Software, 1998). The number of cointegrating relation(s) is significant at the 95% confidence level. (2) See *Log Lines Statistical Yearbook* (Arbor-Pacific Forestry Services Inc., 1998) for definitions of domestic and export log grades, and price regions. Similarly, see *Forest Product Price and Market Statistics Yearbook* (Random Lengths Publications, Inc., 1995–99) for information on lumber grades.

For the two export sawlog price series, all specifications consistently gave rise to zero cointegrating relation. Considering the high price levels and dramatic fluctuations caused by public harvest reductions in the early 90s, this outcome is not surprising. Also, the choice of monthly instead of quarterly or yearly data may have made it more restrictive for price series to be found cointegrated. Therefore, it can be summarized that domestic lumber markets are integrated with a probability higher than domestic log markets, but export log markets are not integrated. Even for those integrated markets, it may not be true that the underlying price series have the same power to predict each other. We will see this point below.

Since the domestic lumber and log markets are likely integrated, causality tests between the corresponding price pairs should be done in a VEC system. Results from the weak exogeneity test are listed in Table 2. We found that domestic log prices for Region 1 significantly cause those for Region 3, while prices for Region 3 fail to cause those for Region 1. Put it another way, Region 1 leads Region 3 in the price formation process. Of the two lumber grades, the 2×6 #2 and Better dimension lumber leads the 2×4–8′ stud in setting prices. Our next task was to examine the causal relationships among domestic log and lumber price series. Again, Table 2 shows that both lumber price series are weakly exogenous to the movements of the domestic log price series for Regions 1 and 3, suggesting that lumber price series lead the domestic log price formation.

We then considered causality relationships related to the two export log price series in two ways. First, since they are not cointegrated, the normal Granger-causality test applies between them. Results in Table 3 suggest that prices for Region 1 tend to Granger-cause the movements of prices for Region 3. Second, we used the weak exogeneity test in a multivariate context including export log price series as well as domestic log or lumber price series. Our results in Table 2 reveal that export log prices for Region 1 lead domestic log prices for both Regions 1 and 3, and export log prices in Region 3 lead domestic log prices for both Regions 1 and 3. Similarly, export log price series in the two Regions lead the stud and dimension lumber prices as well. To sum up, domestic log price movements in the PNW follow the movements of lumber prices, which are in turn led by export log prices.

Table 2
Results of weak exogeneity tests of forest products markets for Douglas fir in the Pacific Northwest

| Price series | Weak exogeneity | Results |
|---|---|---|
| *Lumber price series* | | |
| 2×6 #2 and Better dimension | 0.07 (0.787) | Leading price |
| 2×4 stud | 6.38 (0.011)* | Following price |
| *Domestic sawlog price series* | | |
| Region 1 | 0.30 (0.586) | Leading price |
| Region 3 | 8.70 (0.003)** | Following price |
| *Export and domestic sawlog price series* | | |
| Region 1 (export) | 3.74 (0.053) | Leading price |
| Region 3 (export) | 2.71 (0.100) | Leading price |
| Region 1 (domestic) | 8.37 (0.004)** | Following price |
| Region 3 (domestic) | 5.27 (0.022)* | Following price |
| *Lumber and domestic sawlog price series* | | |
| 2×4 stud | 0.003 (0.955) | Leading price |
| 2×6 #2 and Better dimension | 2.51 (0.113) | Leading price |
| Region 1 (domestic) | 6.60 (0.010)* | Following price |
| Region 3 (domestic) | 5.69 (0.017)* | Following price |
| *Lumber and export sawlog price series* | | |
| 2×4 stud | 15.59 (0.000)** | Following price |
| 2×6 #2 and Better dimension | 12.46 (0.002)** | Following price |
| Region 1 (export) | 1.20 (0.549) | Leading price |
| Region 3 (export) | 5.61 (0.060) | Leading price |

(1) Test for weak exogeneity is based on the null hypothesis that $H_0$:$\alpha_i = 0$. (2) ** and * denote rejection of the null hypothesis at 1 and 5% significant level, and values in parentheses are *P*-values.

Out of curiosity, we also ran the normal Granger-causality test for all the combinations involving the two cointegrated pairs of lumber and log price series. Results are reported in Table 3. Surprisingly, these results are fairly consistent with what we have obtained earlier. Namely, domestic log prices in Region 1 Granger-cause those in Region 3, but not the other way around. Similarly, prices of the 2×4–8′ stud do not Granger-cause those of the 2×6 #2 and Better dimension lumber, whereas dimension lumber prices have power to influence stud prices. Among the relationships of lumber and domestic log markets, although log prices Granger-cause lumber prices based on some of the tests, the causality is generally weak. However, lumber prices, whether stud or dimension, strongly Granger-cause domestic log prices. Among relationships of the export log and lumber markets, we found a two-way causality. That is, stud prices have the power to predict export log prices in both Regions 1 and 3, but the same is not true of dimension lumber prices. On the other hand, export log prices in both regions have a high likelihood of leading the price movement of dimension lumber and studs. Once again, considering that export log prices for Region 1 lead those for Region 3 and that dimension lumber prices follow stud prices, we may downplay the significance of the effect that stud prices exert on export log prices. Therefore, although the pair-wise Granger-causality test looks awkward and tedious, and its results seems vague and blunt, the impact of cointegrating relation(s) on its results is not necessarily substantial in the current context.

## 5. Conclusions

In recent years, cointegration and causality tests have been used to examine market relationships of forest products. However, the distinctions between the two tests have not been well recognized and, in most cases, they have been used in separation.

Table 3
Results of Granger-causality tests of forest products markets for Douglas fir in the Pacific Northwest

| Null hypothesis | Lag 1 | | 2 | | 3 | | 4 | |
|---|---|---|---|---|---|---|---|---|
| D3 does not Granger-cause D1 | 2.02 | a | 2.62 | a | 1.95 | a | 3.62 | r |
| D1 does not Granger-cause D3 | 9.23 | r | 6.09 | r | 4.27 | r | 5.16 | r |
| E1 does not Granger-cause D1 | 11.51 | r | 5.94 | r | 8.77 | r | 7.05 | r |
| D1 does not Granger-cause E1 | 1.29 | a | 0.01 | a | 2.14 | a | 1.59 | a |
| E3 does not Granger-cause D1 | 8.25 | r | 4.44 | r | 5.03 | r | 4.51 | r |
| D1 does not Granger-cause E3 | 0.02 | a | 1.01 | a | 3.18 | r | 2.93 | r |
| LS does not Granger-cause D1 | 12.57 | r | 8.59 | r | 13.81 | r | 11.12 | r |
| D1 does not Granger-cause LS | 1.46 | a | 6.39 | r | 6.21 | r | 5.91 | r |
| LD does not Granger-cause D1 | 15.07 | r | 8.69 | r | 7.83 | r | 6.76 | r |
| D1 does not Granger-cause LD | 1.21 | a | 3.33 | r | 3.22 | r | 5.05 | r |
| E1 does not Granger-cause D3 | 13.72 | r | 7.98 | r | 9.45 | r | 8.29 | r |
| D3 does not Granger-cause E1 | 1.96 | a | 0.59 | a | 1.23 | a | 1.68 | a |
| E3 does not Granger-cause D3 | 7.65 | r | 5.19 | r | 4.25 | r | 4.64 | r |
| D3 does not Granger-cause E3 | 0.52 | a | 2.82 | a | 2.09 | a | 1.75 | a |
| LS does not Granger-cause D3 | 22.34 | r | 12.17 | r | 10.69 | r | 7.77 | r |
| D3 does not Granger-cause LS | 0.25 | a | 3.14 | r | 4.57 | r | 3.74 | r |
| LD does not Granger-cause D3 | 15.23 | r | 11.19 | r | 8.41 | r | 6.16 | r |
| D3 does not Granger-cause LD | 1.01 | a | 2.43 | a | 3.08 | r | 2.92 | r |
| E3 does not Granger-cause E1 | 0.13 | a | 1.25 | a | 1.03 | a | 2.71 | r |
| E1 does not Granger-cause E3 | 12.03 | r | 25.88 | r | 15.88 | r | 13.72 | r |
| LS does not Granger-cause E1 | 4.61 | r | 4.78 | r | 4.69 | r | 3.71 | r |
| E1 does not Granger-cause LS | 1.65 | a | 4.71 | r | 4.17 | r | 6.11 | r |
| LD does not Granger-cause E1 | 0.59 | a | 2.85 | a | 2.66 | a | 2.26 | a |
| E1 does not Granger-cause LD | 7.21 | r | 9.71 | r | 6.01 | r | 10.71 | r |
| LS does not Granger-cause E3 | 8.07 | r | 5.23 | r | 6.01 | r | 4.67 | r |
| E3 does not Granger-cause LS | 1.08 | a | 4.49 | r | 5.37 | r | 4.68 | r |
| LD does not Granger-cause E3 | 0.58 | a | 2.01 | a | 3.13 | r | 2.34 | a |
| E3 does not Granger-cause LD | 7.71 | r | 6.66 | r | 6.46 | r | 5.79 | r |
| LD does not Granger-cause LS | 0.26 | a | 7.15 | r | 4.81 | r | 3.41 | r |
| LS does not Granger-cause LD | 1.06 | a | 0.54 | a | 0.31 | a | 0.34 | a |

(1) D1 and D3 represent domestic sawlog prices in Regions 1 and 3, E1 and E3 represent export sawlog prices in Regions 1 and 3, and LS and LD represent 2×4 stud and 2×6 #2 and Better dimension lumber prices; (2) 'a' and 'r' indicate acceptance or rejection of the null hypothesis; (3) $F$ statistic at the 95% significance level is 3.91, 3.09, 2.69, and 2.46 for lag length of 1, 2, 3, and 4.

We argue that because of their different foci, the two tests can well complement each other. Therefore, using them in combination will provide richer empirical results. That argument gave us the initial motivation to conduct this study.

Of the six markets in the PNW we examined, the two domestic log markets and the two lumber markets are, respectively, integrated, while the two export log markets are not, nor is any cross-grade combination. In conjunction with cointegration restrictions, our causality and weak exogeneity tests further show that the export and domestic log price series for Region 1 lead those in Region 3, and the stud price series follow the dimension lumber price series in the process of price formation. Further, export log and lumber prices lead the movements of domestic log prices, and the movement of domestic lumber prices follows export log prices. Thus, it seems clear that export prices lead the movements of the lumber and domestic log prices, and export log prices in Region 1 lead the price formation process in all the markets—lumber, and domestic and export log ones.

These results have significant implications for understanding and dealing with market behavior and price forecasting. Since their prices move in a similar fashion, integrated markets can be aggregated into a single market (Hamilton, 1994; Johansen, 1995), which makes aggregate market analysis feasible, particularly for long-run equilibrium study. At regional, national, or international levels, aggregation is often inevitable. However, aggregation is more appropriate if we know the markets under consideration are integrated. Aggregation imposed on markets that are not integrated could alter the outcome of market analysis. In our case, the domestic log and lumber markets can be aggregated with confidence, but aggregation of the two export log markets seems problematic. On the other hand, it is equally problematic to say that, if markets are not integrated, then there must be opportunities for arbitrage, since lack of price cointegration is a necessary but not sufficient condition for arbitrage.

Results of our causality tests suggest that prices in the Region 1 export market move first, followed by those in Region 3, and then their movements are transmitted to the domestic lumber and log markets. These relationships of market precedence significantly illustrate the price formation process for forest products in the PNW. Since the Seattle Customs District is by far the largest in the PNW (Warren, 1997), it naturally sets export prices for the region. In addition, industry sources have long claimed that the prospect of the regional domestic log market hinges on the outlook of the export log market (Lynn, 2000; Random Lengths, Inc., 1998). Our analysis has provided supportive evidence for this claim. In a similar manner, we saw that all lumber prices affect domestic log prices. So, lumber prices can be used to forecast log prices, whereas log prices can rarely be used to predict lumber prices. This result may also put in doubt the assertion that the higher lumber prices in the early 90s were partially driven up by the log price hikes induced by the spotted owl debate and ensuing federal policy changes (Random Lengths, Inc., 1998). Moreover, it validates the hypothesis of derived demand, which postulates that, as demand for logs is derived from demand for lumber, lumber prices play a key role in determining log prices (Robinson, 1987).

In addition to indicating which variable(s) can be used to forecast a specific price series of interest, the test statistic of Granger causality also showed at what lag(s) the dependent variable may be most highly associated with the independent variable(s). With this knowledge, it becomes easier to specify and estimate a price forecasting model and obtain better results. It should be emphasized, though, that in the presence of extensive cointegrating relations among the price series considered, the normal Granger causality procedure should be replaced by the weak exogeneity procedure that incorporates the cointegrating restriction(s).

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