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ACCOUNTING FOR EXOGENOUS SHOCKS IN DETERMINING SOUTHEASTERN U.S. TIMBER MARKETS

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Abstract. Cohesive markets for pine sawtimber and pine pulpwood were identified across regions in the southeastern United States. Data were provided by TimberMart-South (TMS) and included 40 years of quarterly prices across 22 regions in 11 states. Markets were determined using the law of one-price cointegration while accounting for endogenous structural breaks. The resulting markets for pine pulpwood resulted in one major market spanning from northern Georgia to southern Texas (seven TMS regions) and four markets made up of two TMS regions. Pine sawtimber markets could be interpreted as nine minimarkets or six markets that were driven by three independent markets made up of the largest mill-capacity regions. Compared to earlier studies, the markets were more fractured. The timing of the endogenous breaks was consistent across regions.

Keywords: Causality; cointegration; forest markets; shocks; time series.

1 Introduction

Timber production in the southeastern United States (Southeast) varies greatly throughout the region because of climate, ease of transportation, geography, proximity to ports, soil conditions, and many other factors (Porter et al., 2014). This suggests that even for a relatively homogenous product, the market might be segmented. The law of one price (LOP) posits that prices for a commodity in a free market should converge over the long term after taking into account transportation costs, transaction costs, and other cost differences among regions (Uri and Boyd, 1990). Moreover, an exogenous price shock, such as a natural disaster, should converge back to a single price over the long term. Measuring the presence of dispersions between subregions can determine local spatial market equilibria. By identifying which regions effectively compete as a single market, buyers and sellers can better forecast which shocks from neighboring regions will most directly affect them, growers can better determine a fair/market selling price, and producers are better able to plan the location of mills ensuring a stable supply. Morales et al. (2015) tested the LOP in global pulpwood markets and while several long-term relationships were discovered, the LOP generally did not hold. Although the LOP has been applied to many commodities and other timber products and markets, the Southeast provides a unique opportunity to compare a large number of subregions that face similar macroeconomic conditions while maintaining unique relationships with each other. TimberMart-South (TMS) data contain nearly 40 years of quarterly prices from 11 states in the Southeast that are well suited for identifying such markets. This period of time encompasses distinct structural breaks in the market that significantly affected market behavior across the region.

Several previous studies used TMS data to identify unique markets. Yin et al. (2002) examined pine sawtimber (PST) and pine pulpwood (PP) prices using appropriately specified pairwise Dickey-Fuller tests for cointegration. Grouped subregions were further verified with Johansen tests. They rejected the presence of the LOP throughout the entire region but did find evidence of subregions that acted as unified markets. Although their market definitions were restricted to geographically cohesive groups, they found evidence of cointegration between geographically distant subregions. Bingham et al. (2003) considered outside policy factors, such as the reduction of timber harvesting on Federal land in 1998, as an exogenous shock. Their results suggested that price shocks are more quickly disseminated across the coastal areas of the Southeast, thus creating one large market. They found evidence of two separate interior markets in the northern and western parts of the region, although those were not defined explicitly.

Zhou and Buongiorno (2006) created a space-time autoregressive moving average model to which they applied impulse shocks. Price shocks were neither statistically nor economically significant past the second-order neighbor and took at most a year to disperse. Rather than defining separate submarkets, each TMS region was treated as the center of its own submarket that overlapped with all the other submarkets. Hood and Dorfman (2015) analyzed the dynamics of the TMS stumpage regions with a time-varying smooth transition autoregressive model. Housing starts were used as an outside indicator variable. They found that all the markets were linked at the peak of demand because of the housing boom. Markets tended to segment more as the demand fell.

Previous studies may not fully account for large exogenous structural changes that affect an entire region. These may not affect individual markets simultaneously or to the same degree. Such changes may include the Staggers Rail Act deregulation in 1980, changes in harvests on Federal lands, Environmental Protection Agency mandates, pest infestations, changing global demands, the digital revolution, and other macroeconomic conditions. Structural changes may accelerate or counteract current trends. Even temporary shocks can cause realignment of the opening and closing of mills thus changing the inherent lumpiness of the way price may be transmitted. Analyses that overlook such changes may lead to spurious relationships between regions.

Using unit root tests that take into account structural breaks allows for a significant increase in the power of the tests and prevents spurious correlation. Moreover, two price series differenced on one another form a new price series that may be tested for structural breaks. The presence of a structural break in this case is referred to as a co-break when the direction of the influence is known (Perron, 2006). By using methods that find the break endogenously, more information is revealed about the behavior of the market over time (Glynn et al., 2007). Applying the econometric results to practical knowledge of the timber markets in the region over time allows for a more accurate definition of the submarkets. Determining the timing of the break, even when not statistically significant, reveals information about the dynamics of the time series data and if previous markets may have been overstated (Yin et al., 2002). Also, the timing of the breaks can indicate when a structural shift occurred and which regions were affected.

The structural breaks affecting the southeastern timber markets represent much more than price shocks. The timber industry is constantly changing in terms of products produced and markets supplied. For example, newsprint demand has declined as bioenergy markets have accelerated in Europe. The structural shifts examined in this chapter are more than the result of normal business cycles and changes in plant capacity utilization. These shocks can cause mills to relocate and can also cause changes in the demand for specific types of timber products. The unique characteristics of timber production, including an unusually long planning horizon for raw inputs, may make shocks to the market more persistent.

2 Methods

Cointegration analysis allows evaluation of whether or not markets follow the LOP and behave as one market. Most prices exhibit nonstationary behavior over time. If their first-differences are stationary, they must be integrated of order 1 [I(1)] (Takayama and Judge, 1964). If a combination of two prices can be expressed as a time series that is stationary [I(0)], the two are said to be cointegrated. This implies that price changes in two spatially separated markets are perfectly transmitted over time, adjusting for exogenous factors such as differences in transaction costs.

The augmented Dickey-Fuller (ADF) test is the standard for determining stationarity (Dickey and Fuller, 1979). Assuming the true model can be represented as a random walk with drift $y_t = \alpha + y_{t-1} + u_t$, where α is the constant drift term, y_{t-1} represents the previous period, and u_t is an independent stochastic error term. By including a time trend and differencing, the standard Dickey-Fuller model can be tested $\Delta y_t = \alpha + \rho y_{t-1} + \delta_t + u_t$ with ordinary least squares. The null hypothesis suggests $\rho = 0$, indicating the time series is nonstationary and contains at least one unit root. It is possible for either α or δ to be equal to zero, and their significance in the equations needs to be determined in order to be properly specified (Hamilton, 1994). The former implies drift, whereas the latter implies a deterministic trend. ADF improves on the standard model by addressing serial correlation by including lag terms of the differenced time series. In this paper, the ADF with both α or δ, as well as an ADF that includes a trend, are used only if it is statistically significant for an individual data series. Because the prices never start at zero, the possibility of the intercept being nonsignificant is dismissed.

Results of the test are sensitive to the number of lags which need to be determined on an individual series' basis (Cheung and Lai, 1995). If the lag number is too small, serial correlation will remain and bias the test. If the number is too large, the test will lose power. The Akaike information criterion (AIC; Akaike, 1973) is represented as $AIC = 2k - 2\ln(L)$, where k is the number of estimated parameters and L is the maximum of the likelihood function. The Schwarz information criterion (BIC; Schwarz, 1978) is represented as $BIC = -2\ln(L) + k\ln(n)$, where n is the number of observations, k is the number of free parameters, and L is the likelihood function. Both reward the goodness-of-fit while penalizing over-parametrization, although BIC is stricter and has better asymptotic properties.

An alternative to the ADF is the Phillips-Perron test (Phillips and Perron, 1988) which also aims to account for serial correlation and heteroscedasticity in the standard Dickey-Fuller test. Rather than lags, it uses a nonparametric approach and adjusts the estimated variance. Compared to ADF, it has the advantage of not needing to specify the number of lags and being more robust to different forms of heteroscedasticity. However, it is more prone to type I errors (rejecting the null hypothesis when true). Due to its structure, the Phillips-Perron test does not allow trend without drift.

The most robust method implemented will be the ADF generalized least squares (ADF-GLS) test as formulated and developed by Elliot et al. (1996). It uses generalized least squares in place of ordinary least squares in the standard ADF. The advantage of the test is a significant improvement in power. Otherwise, the test is similar to the ADF, but on GLS-detrended data. The test significantly improves on Phillips-Perron and ADF in most cases.

Cointegration was tested with pairwise comparisons between stationary data. The resulting array of significant pairings is used to group regions that are cointegrated with all other regions within a group. The standard differencing method was used for the initial pairings. The differencing approach requires subtracting one time series of prices from another. The subsequent series can be tested for unit roots using ADF. Let p_{it} be price $_i$ in time $_t$, and Z_t be price series difference. Given that $Z_t = p_{1_t} - p_{2_t}$ is a linear combination of two I(1) processes, if the Z_t is I(0), the two price series can be said to be stationary. Compared to regressing one price on the other and testing the resulting residual for unit root, the differencing method has two advantages. First, it is symmetric and switching the order of the two prices does not affect the results. Second, it avoids the simultaneity problem which suggests that both markets could be influenced by the same exogenous factors such as macroeconomic policies or outside information (Engle and Yoo, 1987).

The Johansen method (JH) (Johansen, 1995) tests for cointegration over multiple variables such as potential subregions as a whole. The standard VAR model can be estimated using a vector error correction model (VECM) of the basic form:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \varepsilon_t, \qquad (1)$$

where:

- y is a $(K \times 1)$ vector of I(1) variables (in this case $K = 2$),
- Π is the $(K \times K)$ long-run coefficient matrix,
- Γ_i are the $(K \times K)$ coefficient matrices for every lagged variable,
- ε_t is a $(K \times 1)$ vector of normally distributed errors that are serially uncorrelated.

This model is estimated using maximum likelihood. The rank of Π (denoted as r) can be at most K and is equal to the number of characteristic roots or eigenvalues that are significantly different from zero. If $0 < r < K$, then r represents the number of cointegrating vectors. Full rank would imply that the original series are stationary, whereas $r = 0$ implies there are no linear combinations that are I(0). With n I(1) series, full cointegration would imply rank $r = n - 1$.

There are two primary methods of evaluating the JH, the trace test and the maximum eigenvalue test (Johansen, 1995). The likelihood ratio of the trace test can be expressed as

$$LR = -T \sum_{i=r+1}^{K} ln\left(1 - \hat{\lambda}_i\right), \qquad (2)$$

where:

- T is the number of observations; and
- $\hat{\lambda}_i$ is the estimated eigenvalue.

The null hypothesis is that there are r or less cointegrating vectors. The test begins at 0 rank and works up until the null fails to be rejected. With only two series, there is either rank 1 or 0. Rank 1 implies cointegration. The maximum eigenvalue test is similar but will be omitted as it suffers from the multiple comparisons problem . Moreover, the trace test is more robust to excess kurtosis and skewness (Sjö, 2008).

With the JH, there are five options of models in order of least to most flexible: no deterministic terms, restricted constant, unrestricted constant, restricted trend, and trends for both the cointegrating equation and the difference data. The first and the last of these are highly unlikely, leaving us with the choice of the middle three. Juselius (2006) suggested using the Pantula principle of testing from the most restrictive to the least restrictive. When applicable, this principle was followed beginning with the restricted constant specification.

The Zivot and Andrews (1992) model is used to allow for endogenous structural breaks. The breaks are said to be endogenous, because rather than having a predetermined date chosen by the econometrician directly, the test cycles through every period to pick a potential break point which is most favorable for the null hypothesis. The model can take on three forms. The first allows for a change in intercept (A), the second allows for a change in trend (B), and the third allows for both (C). If the null hypothesis is be expressed as:

$$H_0 : y_t = \alpha + y_{t-1} + u_t. \tag{3}$$

The three models can be represented as H_1:

$$A\Delta y_t = \alpha + \delta_t + \theta DU1_t + \rho y_{t-1} + \sum\nolimits_{i=1}^{p} c_i \Delta y_{t-i} + \varepsilon_t, \tag{4}$$

$$B\Delta y_t = \alpha + \delta_t + \gamma DT1_t + \rho y_{t-1} + \sum\nolimits_{i=1}^{p} c_i \Delta y_{t-i} + \varepsilon_t, \tag{5}$$

$$C\Delta y_t = \alpha + \delta_t + \theta DU1_t + \gamma DT1_t + \rho y_{t-1} + \sum\nolimits_{i=1}^{p} c_i \Delta y_{t-i} + \varepsilon_t, \tag{6}$$

where:

- $DU1_t$ represents a dummy variable that equals 1 for any $t > TB$, where TB is the breakpoint chosen endogenously. This allows for a shift in the intercept.
- $DT1_t$ is equal to $t - TB$ for any $t > TB$, representing a shift in trend.
- Model C allows for both adjustments. This model only allows for one structural break.
- The H_0 of a unit root process without a break is rejected if ρ is statistically significant.
- The breakpoint, TB, is tested sequentially and is chosen where the ADF unit root t-statistic is at a minimum.

Clemente et al. (1998) allow for two endogenously chosen structural breaks, extending the work of Perron and Vogelsang (1992). They consider two different types of breaks. When the breaks belong to additive outliers (AOs), then the change is quick, implying an immediate change in slope with no persistence. The innovative outliers (IOs) imply a more gradual change that perseveres over time. This allows a change in intercept and slope. The null hypothesis implies structural changes with unit root:

$$H_0 : \quad y_t = y_{t-1} + \gamma_1 DTB_{1t} + \gamma_2 DTB_{2t} + \varepsilon_t \tag{7}$$

$$H_1 : \quad y_t = \alpha + y_{t-1} + \theta DU_{1t} + \gamma DTB_{2t} + \varepsilon_t \tag{8}$$

where:

- DTB_{it} is a pulse variable such that $DTB_{it} = 1$ when $t = TB_i + 1$ for $i = 1, 2$.
- DU_{it} represents a dummy variable that equals 1 for any $t > TB_i$ for $i = 1, 2$.

In the case of the IO, the model to be estimated is

$$y_t = \alpha + \rho y_{t-1} + \theta_1 DU_{1t} + \theta_2 DU_{2t} + \gamma_1 DTB_{1t} + \gamma_2 DTB_{2t} + \sum_{i=1}^{p} c_i \Delta y_{t-i} + \varepsilon_t. \tag{9}$$

Afterwards, all break combinations for the minimum value of the pseudo t-ratio for testing if $\rho = 0$ are checked. For the AO, the deterministic part of the model is removed by estimating:

$$y_t = \alpha + \theta_1 DU_{1t} + \theta_2 DU_{2t} + \tilde{y}_t, \tag{10}$$

which allows searching for the minimal t-ratio in

$$\tilde{y}_t = \alpha + \rho \tilde{y}_{t-1} + \sum_{i=1}^{p} \omega_1 DTB_{1t-i} + \sum_{i=1}^{p} \omega_2 DTB_{2t-i} + \sum_{i=1}^{p} c_i \Delta \tilde{y}_{t-i} + \varepsilon_t, \tag{11}$$

converging to the unique distribution presented in their paper.

3 Data

TMS stumpage price data from 11 states were analyzed. Alabama, Arkansas, Florida, Georgia, Louisiana, Mississippi, North Carolina, South Carolina, Tennessee, Texas, and Virginia each had data starting in 4Q1976 through 2Q2016. Data are collected on individual timber sales from reporters in each TMS region (http://www.timbermart-south.com). The data are then checked, aggregated, and compiled by the staff at the Frank W. Norris Foundation (Prestemon and Pye, 2000). Each state is divided into two regions. In this paper, each region will be identified by its two-digit state code followed by 1 or 2 denoting region number (Figure 1). Focus was on the average PST and PP prices.

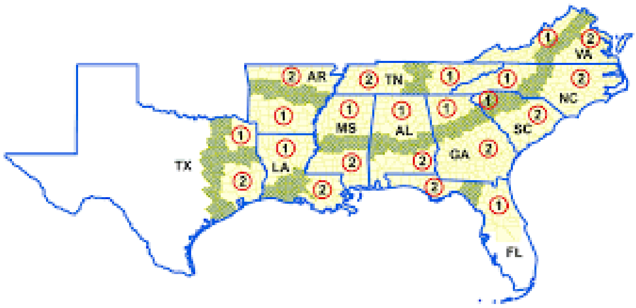

Figure 1: TimberMart-South regions. Each region is designated as two-letter State abbreviation and region 1 or 2.

These were chosen because they are the most consistent in definition and the most complete over time.

Focus was on nominal level prices. Real price data are also analyzed and resulted in similar findings (available from the authors). Usually, the natural log of price is used for cointegration tests. The most common reason is that prices tend to grow exponentially over time. This was not true for either PST or PP as seen in Figures 2 and 3. Although neither exhibited strictly linear behavior, even nominal prices did not exhibit exponential growth. The logarithm of price is also used when the data exhibit great variability, which is not the case in these data. Cointegration tests on log prices imply an interest in percentage change in price rather than the price itself. Given that all regions use the same currency and that changes in price are likely to be equal in level across regions rather than proportional, the log transformation is not necessary.

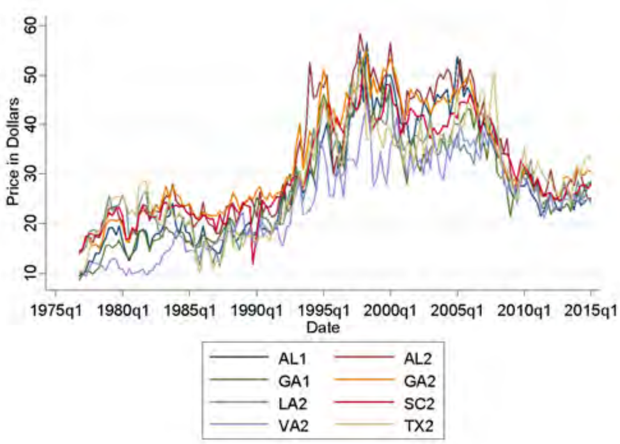

Figure 2: Nominal pine sawtimber stumpage prices for select Timber-Mart South regions. Regions are designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

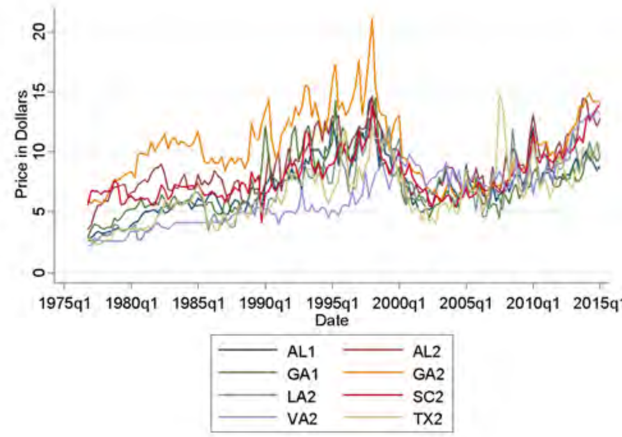

Figure 3: Nominal pine pulpwood prices for select markets over time. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

Two pieces of data are imputed for completeness. Both PST for TN1 in 4Q1997 and PP for AR2 4Q1985 have gaps replaced with the average of the four quarters immediately beforehand and immediately afterwards. TN2 data are dropped completely from PP for having several missing entries in succession. The structural break procedures are sensitive to attempts to impute the missing data. The University of Georgia's Harley Langdale Jr. Center of Forest Business' *Wood Demand Report* provided sawtimber mill capacity data since 1995. Pulpwood mill capacity information since 1980 was supplemented by the *Lockwood Post Directory of Pulp and Paper Mills* as well as the *Pulp and Paper North American Factbook.*

4 Results

No cointegration analysis can be conducted if the initial data are found to be stationary I(0). Table 1 displays the results of 12 different combinations of unit root tests, specifications, and information criteria by column. Each TMS region is tested individually. Only regions that appear to be stationary are shown. The table shows PST and PP testing the nominal price series. Empty columns suggest that all series are found to be nonstationary. Generally, the methods shown are more robust from left to right. The above process was used to determine which of the 22 time series are I(1). If they are of higher or lower order, they are not candidates for cointegration with time series that are I(1). For ADF-GLS, interpolated critical values were used for consistent comparisons (MacKinnon, 1994). The ADF-GLS only allows drift or trend, not both. The number of lagged terms used is determined by either the AIC or BIC. The Phillips-Perron test does not apply lags, but

Table 1: Region tests for unit roots on nominal price data for pine sawtimber and pine pulpwood. [1]

Test	No trend				Trend added			
	AIC		BIC		AIC		BIC	
	PST	PP	PST	PP	PST	PP	PST	PP
ADF	...	AR2* LA1* TN1** TX1* TX2**	...	AR1** AR2* LA1* TN1* TX1** TX2**	...	AR2* LA1* TN1** TX1* TX2**	...	AR1** AR2* LA1* TN1* TX1** TX2**
Phillips–Perron	...	AR2* GA1* TN1* TX1* TX2*	...	AL1* AL2* AR1* AR2* GA1* TN1* TX1* TX2*	NC1** VA1**	AR1* AR2** LA1* LA2** MS2** NC1** NC2** SC1* TN1** TX1** TX2** VA1** VA2*	NC1** VA1**	AL2* AR1** AR2** LA1* LA2** MS2** NC1** NC2** SC1* TN1** TX1** TX2** VA1** VA2*
ADF–GLS	...	...	...	...	...	...	AR2* NC1** TN1* TN2*	...

Note: [1] Tests are more complex from top to bottom and from left to right: ADF = augmented Dickey–Fuller, ADF-GLS = ADF-generalized least squares, AIC = Akaike information criterion, and BIC = Schwarz information criterion. Region is designated as two-letter State abbreviation and region 1 or 2, and regions shown were found to be stationary: * denotes 5% critical value, and ** denotes 1% critical value. PST = pine sawtimber and PP = pine pulpwood.

Newey-West estimation (Newey and West, 1987) uses lags to develop proper standard errors.

Generally, the BIC tests the null hypotheses of non-stationarity and is stricter than AIC. Phillips-Perron is stricter than either of the other two tests. This is consistent with its reputation for type I errors, especially in small samples (Glynn et al., 2007). According to the theory and the literature, the most robust specification is the ADF-GLS model with trend. This is the case where nominal prices perform worse for BIC. Each series is confirmed to be stationary when first-differenced, proving that there are no I(2) series.

Price differenced PST and PP without structural breaks allow for a baseline comparison to the rest of the models. Both AIC and BIC are used to determine the appropriate number of lags (time-differences) in each pairwise test. Then the best grouping was determined, where every TMS region in a proposed market is cointegrated with every other TMS region within the market at a 95% significance level. Every proposed market grouping is tested using JH at 95% certainty level and found to be fully cointegrated. Groupings that could not pass the test are reevaluated.

The definition of a unified market suggests that all members of the market are cointegrated with every other member of the market and they are contiguous. A large market can be broken up by a single territory that prevents a link. Often a TMS region might be cointegrated with two neighbors that are not cointegrated with each other. In this case, a value judgment must be made. The first priority was to match the region with the neighbor that had the most similar list of other cointegrated regions. This tends to favor large markets and leave more single markets. Using this rule, a region that could only be paired with a neighbor that would otherwise be part of a large market would be left stranded. Further judgments took into account outside data. Factors considered included rail and major trucking connections, the presence of large rivers, which would cause bottlenecking at bridges, location of the nearest port, presence of large mills on the border, and relative volumes in production.

Often markets seem to overlap and a decision must be made as to whether to place a TMS region in one market rather than another. Many of the region pairs that pass as cointegrated are unlikely to be in a unified market. It may be possible that LA2 will actually be in a market with VA2 given that they both contain ports, despite the fact that regions differ in climate and environment. This argument would not explain why landlocked AR2 is cointegrated with many regions on the other side of the Southeast U.S. and not with any of the local regions. TN1 is connected with 18 other TMS regions in the BIC case. A higher number of cointegrated pairings does not necessarily represent more credible market connections. A high number of clearly unlikely/spurious pairwise connections force more value judgments and might indicate results that are more arbitrary. Methods that lead to a higher ratio of plausible to implausible pairs are considered more credible.

The pairwise arrays are also created with methods that incorporated structural breaks. This includes Zivot-Andrews (ZA) with intercept, trend, and both intercept and trend breaks and Clemente, Montañés, and Reyes (CMR) methods with one and two breaks for both AO and IO. There is no statistical test to prove definitively which method is superior, although the 2-break methods are considered more powerful. A primary difference between ZA and CMR is the null hypothesis. ZA has a simple unit root, whereas CMR tests against a unit root with structural breaks. The latter is supposed to protect from spurious rejections (Glynn et al., 2007).

Figures 4 and 5 consider the nominal price series for GA1, GA2, NC1 and AL2. AL2 and GA2 are two of the biggest producers in the studied area and neighbors on the coast. GA1 and NC1 are both mountainous regions. Although GA2 is often found to be cointegrated with GA1, NC1 tends to be an outlier cointegrated with very few other regions. The PST prices suggest at least one break. It is not clear from the figures whether the break(s) are better suited to intercept/AO or trend/IO. Overall, PP seemed to exhibit shock(s) in the level of price and PST has more smooth changes. Ideally, one would test for the possibility of more than two breaks. This quickly becomes exponentially more complicated, both technically and theoretically.

The fact that the structural breaks are endogenous allowed us to check whether the predicted break coincides with believable exogenous events and whether these events affected part or an entire region. Table 2 sums up the frequency of structural dates by year. Data are restricted to CMR as both the null and alternative hypotheses contain breaks. The IO is expected to precede AO when accounting for a structural shift at time t since AO is an instant shock, whereas IO is gradual. The change after IO would accelerate to a point that would register as an AO shock.

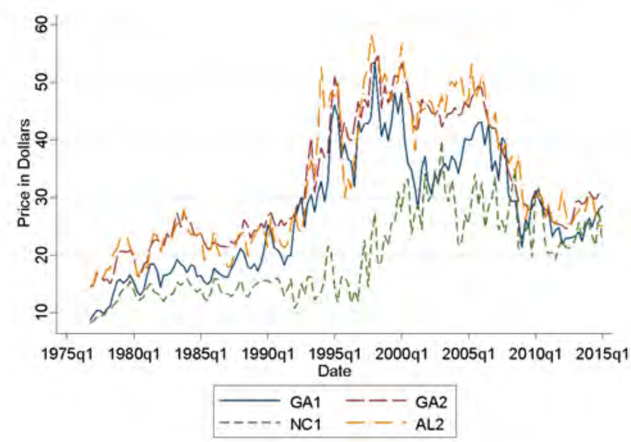

Figure 4: Pine sawtimber time-series samples. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

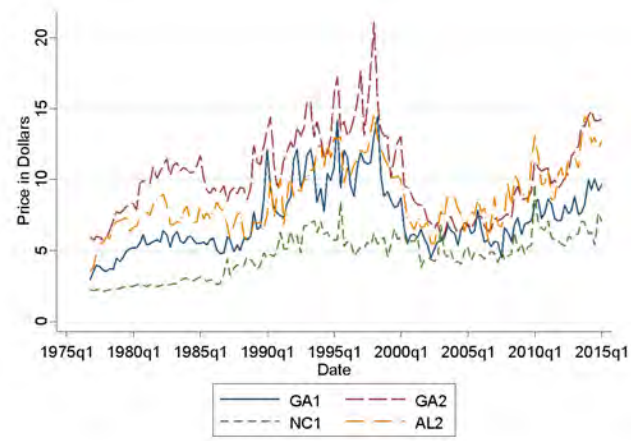

Figure 5: Pine pulpwood time-series samples. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

For PST, Table 2 shows clustering around 1991. When considering two breaks, the breaks are still very bunched with the second break occurring during the start of the housing crisis. The early 1990s was a boom time for Southeast timber with Northwest industry moving to the southeastern United States. IO tends to be slightly more dispersed compared to AO. Reported breaks that edge up to the 5% buffer of the end of the time series may indicate no good candidate for an interior break. CMR must return the most-likely break(s).

Table 2 presents evidence that two shifts have occurred in the southeastern timber market since 1978.

Table 2: Frequency of suggested breaks for pine sawtimber and pine pulpwood using the Clemente, Montañés, and Reyes method and additive outliers or innovative outliers. [1]

Year[2]	1 break				2 breaks			
	PST		PP		PST		PP	
	IO	AO	IO	AO	IO	AO	IO	AO
1986	0	0	0	1	0	0	0	0
1987	0	0	0	0	0	0	0	0
1988	0	0	1	1	0	0	2	0
1989	1	1	3	4	0	0	6	2
1990	1	1	6	2	0	0	5	9
1991	13	0	0	5	2	0	2	1
1992	3	3	3	0	9	2	3	5
1993	2	16	0	0	9	15	0	2
1994	1	0	0	0	1	2	1	0
1995	0	0	0	1	0	2	1	1
1996	1	0	1	1	0	0	2	2
1997	0	0	1	3	2	0	7	3
1998	0	1	0	1	0	1	2	1
1999	0	0	0	0	0	0	2	6
2000	0	0	0	0	0	0	0	0
2001	0	0	0	0	0	0	0	0
2002	0	0	0	0	0	0	0	0
2003	0	0	0	0	0	0	1	2
2004	0	0	0	1	0	0	2	0
2005	0	0	1	0	0	0	0	1
2006	0	0	1	1	6	0	3	1
2007	0	0	0	0	10	7	1	2
2008	0	0	0	0	4	14	1	2
2009	0	0	0	0	0	0	2	3
2010	0	0	0	0	1	1	0	0
2011	0	0	0	0	0	0	0	0
2012	0	0	4	0	0	0	1	1

[1] Note: PST = pine sawtimber, PP = pine pulpwood, AO = additive outlier, and IO = innovative outlier.
[2] No structural breaks were estimated between 1979 and 1986; therefore, those years are not shown.

The ZA tests for structural breaks in the intercept, trend, or both. The CMR tests, in either AO or IO specifications, both test for breaks endogenously. Over two thirds of PST regions show a break in the early 1990s within two years. When the CRM two-break test was applied, 20 of the sawtimber TMS regions show a break between 2006 and 2008 for IO, whereas 21 show a break in 2007 or 2008 for AO.

For PP IO with two breaks, 16 TMS regions showed a break between 1989 and 1992, 13 regions show a break between 1996 and 1999, and seven showed a break between 2006 and 2009. For PP AO with two breaks, 15 show a break between 1990 and 1993, 12 show a break for 1996 through 1999, and nine show a break between 2006 and 2009. Unlike a single shock to a local region, like a hurricane, these structural breaks affect many regions within a relatively short amount of time. Those regions within one market should trend toward equilibrium together.

The suitability of structural break unit root testing on the difference between two stationary processes is identical to that of testing with any other method. The timing of the structural break(s) indicated a permanent or temporary shock to the relationship between two prices. There is no need to assume that all regions in a single market will experience shocks with each other at the same time, although the frequency table does suggest there is bunching. The greater power that comes from these tests provides more assurance that the prices be-

Table 3: Comparison of market identification methods in pine sawtimber.

Method [1]	Stationary regions	Cointegrated pairs	Distinct markets (≥ 2 regions)	1-region markets
Standard differences, BIC, no breaks	0	158	4	0
Standard differences, AIC, no breaks	0	154	7	2
Standard differences, AIC, before 1993	1	140	4	6
Standard differences, AIC, after 1992	1	256	6	6
Zivot–Andrews, break, intercept	5	204	7	2
Zivot–Andrews, break, trend	1	200	7	0
Zivot–Andrews, break, intercept and trend	8	236	5	1
Clemente, Montañés, and Reyes, IO, 1 break	1	158	8	0
Clemente, Montañés, and Reyes, AO, 1 break	4	130	9	2
Clemente, Montañés, and Reyes, IO, 2 breaks	6	80	8	1
Clemente, Montañés, and Reyes, AO, 2 breaks	8	118	5	6

[1] Note: BIC = Schwarz information criterion, AIC = Akaike information criterion, AO = additive outliers, and IO = innovative outliers.

tween two markets are indeed correlated and not spurious.

Figures 6 and 7 show four States from the previous tables with one TMS region's price subtracted from the other. Given the evidence of price breaks in the individual States, major breaks in the early 1990s and last 2000s for the markets suggest an IO approach over an AO approach. This graph also supports the theory that markets tend to diverge during expansion and converge during contraction in industry cycles (Hood and Dorfman, 2015). An IO interpretation is also possible in pulpwood, which seems more volatile.

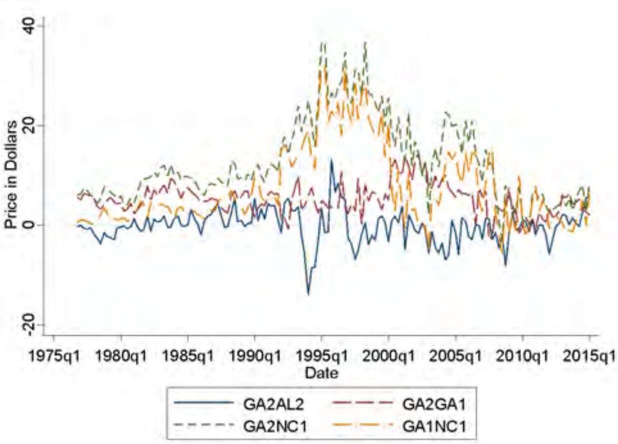

Figure 6: Differenced pine sawtimber stumpage price for select TMS regions (prices in one TMS region subtracted from prices in another TMS region).Regions are designated as two-letter state abbreviation and region 1 or 2; q1 = quarter 1.

Tables 3 and 4 provide a comparison of a sample of considered methods. Stationary regions are the num-

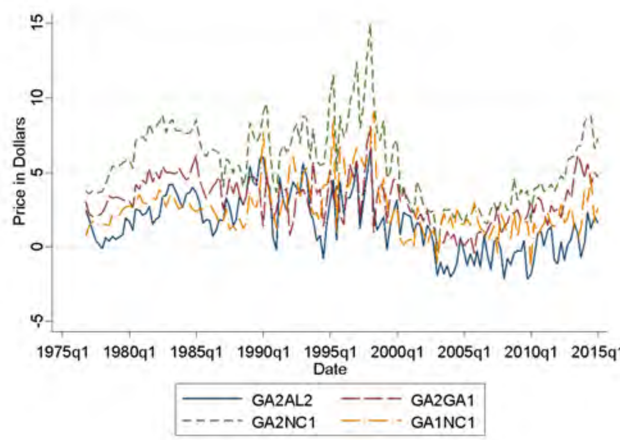

Figure 7: Pine pulpwood cross-region samples. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

ber of TMS regions found stationary when the test was applied on the single region. Given the stronger power of the test, more stationarity is expected than for the standard tests; the regions are predominantly I(0) as shown in Table 1. The stationary regions tend to be isolated with fewer reports and lower production. The two-break methods do exhibit at least one major region each as stationary at the 5% level. This could potentially make them inappropriate for pairwise cointegration testing. However, structural break models are to be an extension of traditional unit root tests, not a replacement. The major regions that are stationary differ between the AO methods and the IO methods.

Cointegrated pairs are the number of the unique pairs that are shown to be stationary. The number of markets is determined as described above, generally first priori-

Table 4: Comparison of market identification methods in pine pulpwood.

Method [1]	Stationary regions	Cointegrated pairs	Distinct markets (≥ 2 regions)	1-region markets
Standard differences, BIC, no breaks	0	162	4	3
Standard differences, AIC, no breaks	0	78	4	8
Standard differences, AIC, before 1993	1	110	4	8
Standard differences, AIC, after 1992	0	68	4	10
Zivot–Andrews, break, intercept	8	106	5	6
Zivot–Andrews, break, trend	3	122	3	9
Zivot–Andrews, break, intercept and trend	7	170	4	6
Clemente, Montañés, and Reyes, IO, 1 break	0	130	6	4
Clemente, Montañés, and Reyes, AO, 1 break	3	134	5	4
Clemente, Montañés, and Reyes, IO, 2 breaks	1	62	6	5
Clemente, Montañés, and Reyes, AO, 2 breaks	4	86	5	5

[1] Note: BIC = Schwarz information criterion, AIC = Akaike information criterion, AO = additive outliers, and IO = innovative outliers.

tizing larger regions, then readjusting to eliminate left-over one-region (single) markets, and finally reevaluating using external real-world consideration such as mill locations, natural barriers and ease of transportation. Alternative interpretations would usually increase the number of markets by splintering them but would not greatly change the number of single regions left over.

A higher number of cointegrated pairings does not necessarily represent more credible market connections. A high number of clearly unlikely/spurious pairwise connections force more value judgments and might indicate more arbitrary results as it increases the possible interpretations. Methods that lead to a higher ratio of plausible to implausible pairs are considered more credible.

The markets for PST and PP in the unit root tests over shorter segments show that simply breaking up the data in smaller time periods is not promising. The data set starts to exhibit much greater small sample issues. Although not apparent here, more series tested stationary as the time periods decreased. The sharp increase in stationarity may be a result of the smaller sample properties (Hosken and Taylor, 2004).

PP markets are more disjointed than sawtimber. Often a majority of regions can be designated as single region markets. The number of cointegrated pairs is also lower. This is probably due to higher proportional transportation cost over revenue for pulpwood. Another factor is the greater lengths mills are willing to take to keep a sawmill supplied over a pulpmill. The CRM with one break stands out as having a comparable amount of cointegrated pairs to sawtimber.

For PST and PP, the CRM IO two-break structural model has the least spurious cointegration array while giving complete and viable maps. IO is more theoretically sound over AO given that persistent shifts are anticipated in markets between regions rather than one off shocks. Using the cointegration array from sawtimber, two different alignments can be derived emphasizing the need for consideration of practical concerns. The first arrangement in Figure 8 tries to incorporate as many regions into markets as possible. Most groupings are composed of timber markets with relatively large mill capacity combined with a market with lesser capacity. The exceptions are FL1 and FL2, which are small capacity markets, AL1 and AL2, which are large capacity markets, and TN1, which is the only single low-capacity market. The issue with this specification is the Arkansas-Louisiana corridor. There is no major freight transportation infrastructure connecting AR1 with LA1, particularly towards the West where most mills are located.

Figure 8: Pine sawtimber stumpage markets based on Clemente, Montañés, and Reyes innovative-outlier 2 breaks optimized for inclusiveness. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

There is a natural alternative specification in Figure 9. Grouping decisions are based on freight transportation patterns. An alternative specification of PST markets contains six PST markets with four single region markets left over. Three of the four single markets have the three largest PST sawmill capacities in the region suggesting their weight has them respond differently to price shocks than the surrounding markets.

Figure 9: Pine sawtimber stumpage markets based on Clemente, Montañés, and Reyes innovative-outlier 2 breaks optimized by freight transportation. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

For PP in Figure 10, contradictory pairings were resolved based on pulp/paper and bioenergy mill locations. Of the five markets, all but one consists of two regions. The seven-region market stretches from TX2 to GA1. The three of the four singletons have mill capacities that tend to be negatively correlated with the rest of the regions and positively correlated with each other. All three have lost more mill capacity proportionally since 1990 than any other region, with the exception of VA2.

Figure 10: Pine pulpwood stumpage markets based on Clemente, Montañés, and Reyes innovative-outlier 2 breaks. Region is designated as two-letter State abbreviation and region 1 or 2; q1 = quarter 1.

5 Discussion and Conclusions

This paper uses TMS price data to determine which TMS regions form a cohesive market. The data are used to determine endogenously the structural break points of each price in each TMS region. A majority of the break points occur during periods that have clear explanations for an outside shock. These shocks were more clustered for PST rather than PP. By taking into account structural breaks, spurious connections are reduced between regions that do not behave as one market but were affected by an exogenous shock to the entire region. PST markets either paired large capacity regions with small capacity regions, or high-capacity regions acted as their own single markets. For PP, the market spanned five states. The three markets that have seen the largest decline in capacity were single markets. The three-time periods around which the endogenous breaks cluster represent a significant restructuring for the industry with persistent effects compared to regular business cycles. The early 1990s structural break stems from the declaration of the northern spotted owl as endangered, resulting in severe restrictions in logging in the northwest United States. The southeastern United States absorbed much of the excess demand as it declined in the Northwest. In the late 1990's the pulpwood market reached a turning point where capacity started falling after consumption peaked in 1994 and exchange rates were unfavorable to exports. The housing crisis, which began in 2006, severely affected the timber industry. Although all the regions were affected, the timing and extent of the reactions differed geographically. For instance, Arkansas, Texas, and Louisiana pulpwood markets were affected by the recession sooner than the rest of the region.

The endogenously determined breaks display a consistent pattern with the proposed structural breaks affecting regions directly. There is no significant bunching at the beginning or the end of the time frame. PST has between 10% and 20% of its restructuring breaks occurring before 1989. PP, which was exhibiting steady growth at the time, had less than 5% breaks prior to 1989. The dispersion of breaks over time is larger for PP than for PST.

The structural changes between regions are expected to be persistent. Compared to the nonstructural baselines, the groupings are more conservative, reflecting the increased power of the tests. The results are less dependent on individual interpretation than previous methods. Six integrated markets were identified along transportation corridors. The three regions with the largest mill capacities are independent. Pine pulpwood has one large unified market ranging from lower Texas to the mountains of Georgia. Three out of the four one-region markets showed the largest proportional downturn in production since the 1990's. In both cases, these groupings survived the restructuring of the industry over the last 40 years.

It may be insightful to compare the nature of the results with those of Hood and Dorfman (2015). Both acknowledge that the market changes over time. Although their STAR model shows ebb and flow of markets quarter to quarter, our study aims to cut through to fundamental relationships that span over transition periods. The markets evaluated in this paper represent those that have persevered through significant positive and negative shocks to the market as well as the advancements in technology and evolution of the global market. The groupings imply fundamental underlying characteristics (geography, forest resources, transportation infrastructure, etc.) that dispel price shocks more proficiently.

There are several limitations of this study that could be addressed and expanded upon with more data. Further analysis into the markets could be conducted with more complete and detailed production and production-capacity data. Calculating a supply curve would be possible with data on both quantity and price. It would be viable to see how production shifts between regions within a unified market given short-term exogenous shocks or region-wide structural shifts. Elasticities by region could be determined. Demand information would allow the modeling of the cohesive market. With detailed information about major mill and plant closings and locations over time, it would be possible to expose patterns along the borders of states and see productions shifted from one border to another, verifying a realignment in markets.

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References

Akaike, H. (1973). Information theory and an extension of the maximum likelihood principle. In *Second International Symposium on Information Theory* (pp. 267–281). Budapest: Akadémiai Kiadó.

Bingham, M. F., Prestemon, J. P., MacNair, D. J., & Abt, R. C. (2003). Market structure in US southern pine roundwood. *Journal of Forest Economics, 9*(2), 97–117. https://doi.org/10.1078/1104-6899-00025

Cheung, Y. W., & Lai, K. S. (1995). Lag order and critical values of the augmented Dickey-uller test. *Journal of Business and Economic Statistics, 13*(3), 277–280. https://doi.org/10.1080/07350015.1995.10524601

Clemente, J., Montañés, A., & Reyes, M. (1998). Testing for a unit root in variables with a double change in the mean. *Economics Letters, 59*(2), 175–182. https://doi.org/10.1016/S0165-1765(98)00052-4

Dickey, D. A., & Fuller, W. A. (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association, 74*(366), 427–431. https://doi.org/10.1080/01621459.1979.10482531

Elliott, G., Rothenberg, T. J., & Stock, J. H. (1996). Efficient tests for an autoregressive unit root. *Econometrica, 64*(4), 813–836. https://doi.org/10.2307/2171846

Engle, R. F., & Yoo, B. S. (1987). Forecasting and testing in co-integrated systems. *Journal of Econometrics, 35*(1), 143–159. https://doi.org/10.1016/0304-4076(87)90085-6

Glynn, J., Perera, N., & Verma, R. (2007). Unit root tests and structural breaks: A survey with applications. *Revista de Métodos Cuantitativos para la Economía y la Empresa, 3*(1), 63–79.

Hamilton, J. D. (1994). *Time Series Analysis.* Princeton, NJ: Princeton University Press.

Hood, H. B., & Dorfman, J. H. (2015). Examining dynamically changing timber market linkages. *American Journal of Agricultural Economics, 97*(5), 1451–1463. https://doi.org/10.1093/ajae/aau151

Hosken, D., & Taylor, C. T. (2004). Discussion of "Using stationarity tests in antitrust market definition." *American Law and Economics Review, 6*(2), 465. https://doi.org/10.1093/aler/ahh008

Johansen, S. (1995). *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models.* Oxford, UK: Oxford University Press.

Juselius, K. (2006). *The Cointegrated VAR Model: Methodology and Applications.* Oxford, UK: Oxford University Press.

MacKinnon, J. G. (1994). Approximate asymptotic distribution functions for unit-root and cointegration tests. *Journal of Business and Economic Statistics, 12*(2), 167–176. https://doi.org/10.1080/07350015.1994.10510005

Morales Olmos, V., & Siry, J. (2015). Global pulpwood markets and the Law of One Price. *Mathematical and Computational Forestry & Natural-Resource Sciences (MCFNS), 7*(1), 16–32. https://mcfns.com/index.php/Journal/article/view/MCFNS7.1_3

Newey, W. K., & West, K. D. (1987). A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica, 55*(3), 703–708. https://doi.org/10.2307/1913610

Porter, E., & Consoletti, W. (Eds.). (2014). *How forestry came to the Southeast: The role of the Society of American Foresters*. Cenveo Publisher Services.

Perron, P. (2006). Dealing with structural breaks. In *Palgrave Handbook of Econometrics* (Vol. 1, pp. 278–352). Basingstoke, UK: Palgrave Macmillan.

Perron, P., & Vogelsang, T. J. (1992). Nonstationarity and level shifts with an application to purchasing power parity. *Journal of Business and Economic Statistics, 10*(3), 301–320. https://doi.org/10.2307/1391544

Phillips, P. C. B., & Perron, P. (1988). Testing for a unit root in time series regression. *Biometrika, 75*(2), 335–346. https://doi.org/10.2307/2336182

Prestemon, J. P., & Pye, J. M. (2000). A technique for merging areas in Timber Mart-South data. *Southern Journal of Applied Forestry, 24*(4), 219–229. https://doi.org/10.1093/sjaf/24.4.219

Schwarz, G. (1978). Estimating the dimension of a model. *Annals of Statistics, 6*(2), 461–464.

Sjö, B. (2008). Testing for unit roots and cointegration. *Nationalekonomiska Institutionen, Linköpings Universitet*. 25 p.

Takayama, T., & Judge, G. G. (1964). Equilibrium among spatially separated markets: A reformulation. *Econometrica, 32*(4), 510–524. https://doi.org/10.2307/1910175

Uri, N. D., & Boyd, R. (1990). Considerations on modeling the market for softwood lumber in the United States. *Forest Science, 36*(3), 680–692. https://doi.org/10.1093/forestscience/36.3.680

Yin, R., Newman, D. H., & Siry, J. (2002). Testing for market integration among southern pine regions. *Journal of Forest Economics, 8*(2), 151–166. https://doi.org/10.1078/1104-6899-00009

Zhou, M., & Buongiorno, J. (2006). Space-time modeling of timber prices. *Journal of Agricultural and Resource Economics, 31*(1), 40–56. https://doi.org/10.22004/ag.econ.10147

Zivot, E., & Andrews, D. W. K. (1992). Further evidence on the Great Crash, the oil-price shock, and the unit-root hypothesis. *Journal of Business and Economic Statistics, 10*(3), 251–270. https://doi.org/10.2307/1391541