Comment and Reply

Informational Efficiency of Markets for Stumpage: Comment.

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In a study of the information efficiency of markets for stumpage in the southern United States, Washburn and Binkley found that time series data on stumpage price, if given on annual or quarterly bases, reveal weak-form information efficiency for stumpage markets. From this, they conclude that "there can be no gain from using past price movements to play the market in timing timber harvests" (p. 403). Specifically, they infer that the reservation price rules derived for optimal forest harvesting by Norstrom, Lohmander, and Brazee and Mendelsohn are not applicable to the management of southern pine timberlands.

This conclusion follows if stumpage prices are nonstationary (martingales). However, from a theoretical perspective, it is not obvious that the price of stumpage will be nonstationary, even if stumpage markets are efficient, i.e., maintain rational expectations equilibrium. Moreover, reexamination of the annual and quarterly stumpage price series that were analyzed in Washburn and Binkley's study show, when each series is analyzed separately, that although nonstationarity cannot be rejected (except in one case), neither can the stationarity conjecture. However, treating quarterly data for different states as panel data, a model with pooled data can be estimated and in this model, nonstationarity is rejected.

Therefore, it may indeed be possible for forest owners to enhance forestry income by adapting the timing of stumpage sales to price movements. A brief theoretical background to this issue is provided in this comment and is followed by an account of an empirical test of the nonstationarity conjecture for stumpage prices in the southern United States.

Theoretical Framework

A stochastic price process P_t over time t is a submartingale with respect to a sequence of information sets I_t if P_t has the following property:

$$\text{(1)} \qquad E(P_{t+1}|I_t) \geq P_t$$

where E is the expectations operator. The process is a martingale if this relation holds with equality. In *ex post* form, the martingale and the submartingale correspond to the random walk (or Brownian motion) process and the random walk with drift, respectively. Weak-form efficiency of a capital market (as defined by Fama) means that no trading rule based on historical prices alone can succeed, on average.

Following Washburn and Binkley, let us assume that the growth rate of a forest stand ("stumpage store") at time t is g_t, and the capital cost and storage (maintenance) cost, expressed as proportions of stumpage price, are r_t and c_t, respectively. Also, assume that the stumpage price is nonnegative. The intertemporal arbitrage equilibrium condition for stumpage is then

$$\text{(2)} \qquad P_t = E[(P_{t+1}|I_t)\exp(g_t - c_t - r_t)].$$

In logarithm form, this is (assuming that all variables are known, except the price in the next period)

$$\text{(3)} \qquad \log(P_t) = \log(E(P_{t+1}|I_t) + g_t - c_t - r_t.$$

Assume that the growth rate of the forest stand is a positive, continuous, monotonously decreasing function $g(a_t)$ of the rotation period of the forest, a_t. Also, assume that it can be postulated that the log of the stumpage price process is a martingale, i.e., that stumpage price follows geometric Brownian motion. Then the arbitrage equilibrium equation can be reduced to

$$\text{(4)} \qquad g(a_t) = r_t + c_t.$$

Thus, the harvesting decision at time t depends on capital cost and storage cost, but is independent of the realization of the stumpage price. In other words, the instantaneous stumpage supply is completely inelastic. For nonnegative stumpage prices and constant capital and storage cost, the supply of stumpage will be uniquely determined by the initial age distribution of trees. If the marginal access cost for timber varies, then timber supply will be elastic with respect to the price of timber, reflecting the variation in the volume of stumpage with nonnegative stumpage price as the timber price varies.

The arbitrage equilibrium (4) is in fact equivalent to the Faustmann-Pressler-Ohlin rule for the optimal rotation period under deterministic conditions (in the one-rotation case, i.e., disregarding the value of bare land). A more general analysis of this case is made

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by Clarke and Reed (1989)[1]. Besides assuming that stumpage price follows geometric Brownian motion, they allow for a stochastic age-dependent growth process. Their analysis confirms the above conclusion that the level of the current stumpage price (if it is positive) is irrelevant to the forest owner in his decision on whether to cut now or later. In the Wicksellian wine-aging case (the one-rotation problem), the optimal stopping (cutting) rule resembles the Wicksellian deterministic solution, except for a modification of the rate of interest. The one-rotation tree should be cut when the expected relative growth rate (the deterministic part of the growth function) is equal to the rate of interest less a variance term.

The latter term is shown to be half the sum of the variances and covariance of the price and growth processes. Clarke and Reed (1989) extend this to the case of a risk-averse forest owner and show that the "usual" risk-cost term will be added to the risk-free rate of interest. The case of the on-going forest, with subsequent rotations, turns out to be more tricky, but they are able to derive a similar modified "Faustmann-Pressler-Ohlin" rule in the case where price is the only stochastic variable. A general feature of these results is that the optimal cutting age is independent of the current stumpage price of timber.

Thus martingale stumpage prices give (modified) "Faustmann-Pressler-Ohlin" management. Returning to equation (3), we see that the opposite implication also holds: If equation (4) is valid, then in arbitrage (rational expectations) equilibrium, the stumpage price will be a martingale.

However, assume that it can be postulated that the stumpage price process is stationary. This means that the elasticity of expected future prices with respect to the current price (cf. Hicks, p. 205, Scheinkman and Schechtman, p. 433) is less than one. For example, assume that price follows this process:

$$\log(p_{t+1}) = \alpha + \beta\log(p_t) + u_t,\ 0 \le \beta \le 1 \quad (5)$$

and

$$E(u_t|I_t) = 0 \quad (6)$$

where α and β are constants.

Arbitrage equilibrium (eq. (2)) now gives the following forest management rule:

$$g(a) = (1 - \beta)\log(p_t) + c_t + r_t - \alpha. \quad (7)$$

Clearly, the rotation age is now also dependent on the current price realization, p_t. Differentiation with respect to p and a yields:

$$\frac{da}{dp} = \frac{(1-\beta)}{g'p} \le 0 \quad (8)$$

where $g' = \delta g/\delta a$. This means that the short-run supply curve has a positive slope (an increase in stumpage price leads to a reduction of the volume of the growing stock) even if the marginal access cost is constant.

The implications for optimal forest management of stationary prices have been worked out by Lohmander; related works are Norstrom, Brazee and Mendelson, and Haight. Under stationary prices, the optimal cutting rule can be expressed as a reservation price. Therefore, the harvesting decision (and hence the short-term supply of timber) will depend on the current stumpage price. The nature of the stumpage price process is thus important to forest managers (and to forest economists).

From a theoretical perspective, one would want to know whether a stationary stumpage price process can be upheld in an economy where forest owners and timber purchasers cannot be constantly fooled, i.e., in an economy where agents form rational expectations. For an exchange economy (i.e., no production is undertaken) with risk-neutral agents, Lucas has shown that the price process will be a martingale. However, with risk aversion, this is no longer generally true. Moreover, in a production economy, with risk-neutral agents, the rational equilibrium price process may be stationary. This was first shown by Samuelson for a simple (wheat) production economy. In his model, production in each period is stochastic and the produced good can be stored from one period to another. Scheinkman and Schechtman extended this model by allowing the possibility of raising production by increase in "effort." Scheinkman and Schechtman show that the (point) elasticity of price expectations with respect to the current price is less than one and may be zero for sufficiently high prices. They give the following explanation: "The intuition behind this is that, if today's price goes up, the increase in tomorrow's prices is caused solely by the decrease in storage, since the effort effect causes a decrease in future prices. Thus, if storage levels are already low, further price rises have little effect on future prices" (p. 434).

Therefore, when Washburn and Binkley (p. 394) claim that the reservation price approaches found in Lohmander and others, relies on "market inefficiency" they seem to overstate their case. Inefficient markets can perhaps produce a stationary price process, but so can an efficient market.

Testing for Unit Root

As is recognized by Washburn and Binkley, the validity of the reservation price approach is an empirical issue. A natural point of departure for the empirical analysis presented here would be equations 5 and 6. The martingale conjecture implies a unit β-coefficient (unit root). If β is equal to one and equation 6 holds, in a regression of the residuals on any subset of the information set I_t, then clearly the reservation price approach can be dismissed. If, however, β is below one, then the price process is stationary.

[1] Related works are Clarke and Reed (1990) and Morck, Schwartz, and Stangeland.

In a way, this is the approach used by Washburn and Binkley.[2] Unfortunately, they do not address the stationarity issue directly. Therefore, while their study shows that the conjecture that equation (6) is fulfilled cannot be rejected for annual and quarterly data, we do not know whether this is so for a stationary or a nonstationary process.[3] This problem is, however, easily resolved. Table 1 lists results from regressions of equation (5) on the same data set for quarterly and annual prices as that used by Washburn and Binkley.[4] The results that are reported are the estimates for autoregression coefficient (β), the Dickey-Fuller statistic (DF),[5] and the Durbin-Watson statistic (DW).

Beginning with the last column, the Durbin-Watson tests support Washburn and Binkley's finding that evidence of serial correlation in the residuals (in the first lag) is absent. Thus the DW statistic does not reject the conjecture that equation (6) is valid. In the first column, we see that all estimated β-coefficients are below one. A one-variable ANOVA test, using the estimated β-coefficients and the corresponding standard deviations, does not falsify the conjecture that all quarterly stumpage price series have the same β-coefficient. The unweighted mean of the estimated β-coefficient for quarterly data is 0.90, for annual data, 0.89.

The crucial question is whether the autoregressive coefficients are below one. The DF-statistic is the "t-value" corresponding to the test $\beta = 1$. If it had been known that β is less than one (thus, the test had been superfluous), then the critical value for this test would have been the standard one. However, if the series is nonstationary, then the critical value at, for example, the 5% level is raised from the standard approximately two to approximately three (Dickey and Fuller, Fuller).

As shown by the DF statistic, all estimated β-coefficients, except one, are between one and two standard deviations below unity. The exception is quarterly data for Oklahoma, exhibiting a DF value that

[2] To be meaningful, P_t in equation (5) should be interpreted as the relative price of stumpage, i.e., the real price. Washburn and Binkley make their analysis with nominal prices and regress them on the inflation rate (and stock market return, respectively). A zero intercept and a unit coefficient for the inflation rate imply unit root in our equation (5). Given the specific statistical problems in tests for unit root, the direct approach taken here, i.e., to regress equation (5) directly on deflated prices, seems preferable.

Washburn and Binkley interpret the inflation rate as a measure of the opportunity cost of capital. They also use the stock market return for this purpose. The estimated regression coefficients of the stock market return variable, as well as that of the rate of inflation variable, are in most cases not significant. This either indicates that (nominal) stumpage prices are insensitive to the opportunity cost of capital, or that the wrong capital cost variable is used; which in both cases implies that the test model is misspecified.

[3] For monthly data, Washburn and Binkley find negative correlation at the first lag. They therefore conclude that there may be some scope for the reservation—price approach, if a timber sale can be completed within a month.

[4] Quartal data have been updated up to the first quarter of 1991. Annual data were updated to 1990 (for Louisiana: 1988). Prices have been deflated with the U.S. Consumer Price Index (International Financial Statistics, various issues).

[5] This coefficient is defined as $(\hat{\beta} - 1) / \hat{\sigma}_\beta$ where $\hat{\beta}$ is the estimated autoregression coefficient and $\hat{s}_\beta$ is the estimated standard deviation. See Dickey and Fuller (1979). Critical values for the statistic "$\hat{\tau}_\mu$" are listed in table 8.5.2 in Fuller (1976), p. 373.

Table 1. Estimates of the Autoregressive Equation (Eq. 5)

Price series:	β	DF	DW
Quarterly (Timber Mart South)			
Alabama	0.92	−1.37	1.77
Arkansas	0.95	−1.02	1.77
Florida	0.93	−1.08	1.38
Georgia	0.90	−1.56	1.61
Kentucky	0.88	−1.18	1.75
Louisiana	0.95	−1.16	1.86
Mississippi	0.93	−1.31	1.87
North Carolina	0.94	−1.19	1.48
Oklahoma	0.68	−3.36	1.72
South Carolina	0.87	−1.83	2.23
Tennessee	0.87	−1.76	2.06
Texas	0.95	−1.04	1.64
Virginia	0.84	−1.97	1.92
Quarterly (state data)			
Louisiana	0.97	−1.36	1.55
Mississippi	0.93	−1.08	1.90
Annual (state data)			
Arkansas	0.86	−1.81	1.82
Louisiana	0.90	−1.32	1.64
South Carolina	0.90	−1.28	2.06

rejects unit root at the 2.5% significance level. For all other regions, however, the unit-root conjecture cannot be rejected. On the other hand, if we instead choose to test the hypothesis that the price series are stationary, i.e., $\beta < 1$, then clearly this cannot be rejected either, because all the estimated β-coefficients are less than one.

Thus, analyzing each series independently, we have an indeterminate case. The statistical instruments for making the distinction between a stationary process with an autoregressive coefficient that is close to one, and a nonstationary process are simply not sharp enough. However, it is possible to enhance the precision of the test by pooling data. The data set can be regarded as panel data, containing state-specific measures of a single price series, i.e., the "aggregate" Southern Pine sawtimber stumpage price. Therefore, it is possible to combine the data from different states.

Assume that the stumpage price $P_{i,t}$ in state i is a proportion, m_i, of the "aggregate" stumpage price P_t:

$$P_{i,t} = m_i P_t. \qquad (9)$$

Taking logarithms of both sides, we get $\log(P_t) = \log(P_{i,t}) - \log(m_i)$. Inserting this expression into equation (5) gives the following equation:

$$\log(P_{i,t+1}) = \gamma_i + \beta \log(P_{i,t}) + \epsilon_t \qquad (10)$$

where γ_i is a state-specific parameter $[\gamma_i = \alpha + (1 - \beta)\log(m_i)]$.

Table 2 shows the results of estimating equation (10) on quarterly stumpage prices reported from 1977 by Timber Mart South.[6] The estimated coefficients indicate that the nonstationarity conjecture can be rejected.[7] The estimated β-coefficient is 0.93. The results indicate that the price processes in all states except one (Tennessee) have the same intercept.

More unambiguous answers regarding stationarity for individual price series can be given in other cases. For example, based on the DF statistic, the unit root can be rejected for the annual net unit conversion value[8] of roundwood in Sweden from 1955/56–1987/88 (Hultkrantz 1991), and for monthly stumpage prices in the U.S. Pacific Northwest from 1975–1989.[9]

An important implication of stationary stumpage prices would be that the supply responses to changes in nonnegative stumpage prices are consistent with efficient forest management. Several econometric studies in different countries and regions have confirmed that timber supply from private nonindustrial forest owners is positively correlated with the current price of timber. Of specific interest here is a study by Newman, which analyzes the softwood stumpage markets in the southern United States from 1950–1980. Newman estimated responses of pulpwood and solidwood supply and found own price elasticities of pulpwood stumpage and solidwood stumpage to 0.23 and 0.55, respectively.[10] None of these econometric

[6] Alabama, Arkansas, Florida, Georgia, Louisiana, Mississippi, North Carolina, South Carolina, Tennessee, Texas, and Virginia. The price series for Kentucky and Oklahoma begin later and have therefore not been included.

[7] The small-sample distribution for this specific equation is unknown (for an overview of the small-sample distributions for auxiliary Dickey-Fuller regressions that have been tabulated, see Haldrup and Hylleberg (1991)). However, because the critical value in the standard Dickey-Fuller test at the 1% significance level is −3.44 (−2.57 at the 10% significance level), the estimated DF statistic, −4.29, seems sufficient for rejection of the nonstationarity conjecture.

[8] In Swedish logging, pulpwood and sawtimber are sorted out in the forest. Delivery sales of pulpwood and sawtimber were more common than stumpage sales in the 60s and 70s. Prices are set annually, for the "logging season."

[9] The estimated β-coefficient is 0.90. The DF-value is 2.96, showing significance at the 95%-level. Data are reported in Haynes (1991).

[10] These effects are significant at the 95% and 99% level, respectively.

Table 2. Estimates of the Autoregressive Equation (Equation 5) on pooled quarterly data from 1977 (first quarter)–1991 (first quarter) for Southern Pine Sawtimber Stumpage Prices

Item	Coefficient	t−values
α	0.097	(3.45)
$\hat{\beta}$	0.93	(54.95)
γ_i:		
Alabama	0.0046	(0.23)
Arkansas	−0.0045	(−0.23)
Florida (the no−dummy case)		
Georgia	0.0080	(0.40)
Louisiana	−0.0016	(−0.08)
Mississippi	0.0017	(0.09)
North Carolina	−0.0048	(−0.24)
South Carolina	0.0042	(0.21)
Tennessee	−0.0503	(−2.14)
Texas	0.0019	(0.10)
Virginia	−0.0252	(−1.19)
DF	−4.29	
$\bar{R}^2$	0.91	

studies investigates, however, whether the price response is the result of "access cost" variation or of dynamic reservation price behavior. Hultkrantz (1987) compares an estimated access cost curve for roundwood in northern Sweden with price elasticities estimated on time-series data, and he finds these broadly consistent with one another. However, this is a tentative suggestion and is not founded on an explicit test.

In examining the empirical basis for the different theoretical prescriptions for forest management under price uncertainty, Washburn and Binkley have made an important contribution. However, the evidence does not make it possible to discard the reservation price approach.

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