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Stumpage Prices in Sweden 1909-2011:

Testing for Non-Stationarity*

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Abstract

The price of timber stumpage is one of the few natural-resource rents that can be directly observed as a market price. Rules for optimal timber harvesting under uncertainty have been found to depend on whether the timber rent price is non-stationary or stationary. In this study we extend previous research by Hultkrantz (1995) that tested for unit-root with an exogenous break point in Swedish stumpage prices from 1909-1990, employing data up to 2011, hence for 103 years, and unit-root tests with an endogenously selected break point. We find support for a structural level break at the end of WW2 and that non-stationarity can be rejected. We show that this is a robust conclusion.

JEL classification: Q23

Keywords: Roundwood, timber, natural-resource rents, unit root

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1. Introduction

Stumpage price is the price for a right to harvest a unit of standing timber for which the buyer will bear the cost of felling and transportation. It represents timber rent and is one of the few natural-resource rents that can be directly observed as a market price. Much attention is paid in the natural-resource and forestry management literature to the features of the dynamic process of rents, especially to whether it is non-stationary, but due to low statistical power empirical investigation requires longer time-series than what usually can be found. In this study we extend previous research by Hultkrantz (1995) that tested for unit-root with an exogenous break point in Swedish stumpage prices from 1909-1990, employing data up to 2011, hence for 102 years, and unit-root tests with endogenously determined breaks.

The natural-resource rent, or scarcity rent, has a central role in natural-resource economic theory, building on traditions emerging from works by Ricardo, Jevons, Wiksell, Fischer, Hotelling and others. In forest economic theory, the properties of the dynamic process of timber rent has two important implications. First, if it includes a positive trend component, this provides an additional source of value growth, and therefore enhances the profitability of silvicultural effort and prolongs optimal rotation periods. Second, the decision criterion for choosing the time to cut (the stopping rule) depends on the order of integration of the price process. For instance, if the net price follows a random walk (geometric Brownian motion) a myopic optimal stopping rule will not depend on the current price, while if it is stationary (mean reverting) it does so.

Stumpage price data is more difficult to gather than information about roundwood prices that often is publicly announced. However, stumpage price information can be held from forest-company records. Based on data collected from 1909 to 1955 by Streyffert (1960) and subsequently by the Swedish National Forest Board, Hultkrantz (1995) analyzed times-series data for Swedish stumpage prices from 1909-1990. The National Forest Board terminated its collection of stumpage price data in 2002, but for this study we have been able to update the time series up to 2011, using data held from a stumpage-broker firm (Rotpostmäklarna AB). Perron (1989) showed that the unit-root tests advanced in the pioneering work by Dickey and Fuller (1979) may be misleading if there has been a structural break in the studied time series and suggested a test for the case when there is one known break point. This Perron test was used by Hultkrantz (1995), who found that after accounting for a structural break at the end of WW2, the null hypothesis of unit root for the Swedish stumpage prices could be rejected. During the more than twenty years that have passed since Perron´s contribution, the literature on unit-root testing on data with possible breaks has progressed, so we will here redo the analysis using a method that endogenously determines if and when there is a possible break. Our main result is that the finding in Hultkrantz (1995) is supported.

The paper is organized in the following way. Next section provides a brief background on how variation over time of timber price (net of costs) is thought to affect timber harvesting, i.e., the timber supply. As this depends on the nature of the price process, this section provides the basic motivation for our study. Section 3 then presents the stumpage-price data for the period 1909-2011 and speculates on the role of structural breaks connected to the major technological shifts in timber harvesting. Section 4 reviews statistical methods that can be used to test for unit root in time series that possibly contain structural breaks and motivates our choice of approach for the test. The test results are presented in section 5. The study is rounded up with a discussion and some brief conclusions in the final section.

2. Timber supply and the timber price process

Classical forest economics derived rules for forest management rules that disregarded uncertainty, most famously the Faustmann-Pressler-Ohlin (FPO) rule for determination of the optimal rotation period over an infinite planning horizon (see Johansson and Löfgren 1985 and Amacher et al. 2009). However, a forest is a capital asset and harvesting is basically a de-investment, i.e., an action that involves substitution between production and consumption in different time periods. It therefore involves a short-term planning consideration on whether now is the right time to sell stumpage. There is therefore need for a so called stopping rule on whether the timber growth process should be halted (Amacher et al. 2009, Ch. 11; Davis and Cairns 2012). The stopping rule makes harvesting contingent on currently available information, which of course could include the current price level. However, drawing on elementary asset pricing theory some analysts have conjectured that if traders on the stumpage market take into account all available information, i.e., if this market is informational efficient, the price determination will give rise to a non-stationary, first-order integrated, stumpage price (Washburn and Binkley 1991). This means that the price is a so called martingale that contains no information on whether price next year will be higher or lower, except for a possible deterministic drift component. Indeed, as reviewed in the textbook by Amacher at al. (2009, p. 296) a large amount of research in forest economics theory has followed this route. For instance, based on this assumption, Clarke and Reed (1989) derived a modified FPO rule for a case with stochastic prices. They showed that with age-dependent growth, stumpage supply will depend on the age-structure of the forest inventory (as supply will be given by the "pre-determined" rotation periods) and the opportunity cost of capital.[1] As long as stumpage prices are non-negative the current level of the price should not affect

[1] Provided that all stands yield a non-negative stumpage value at the age of the optimal rotation. A change of the stumpage price may have an effect on short-run supply by changing the share of the forest-land that is economically accessible.

the decision to sell.[2] This result is consistent with the classic FPO rule[3], but is inconsistent with another fundamental finding in the forest economics literature, namely, that the short-run timber supply is price elastic (see for instance Hultkrantz and Aronsson 1989).

In contrast, if the stumpage price is mean reverting (see e.g. Lohmander 1983, Brazee and Mendelsohn 1988, Davis and Cairns 2012), then the current price conveys a signal to the forest owner on whether to sell or wait. To see this in a simple way, as in Hultkrantz (1993) consider the supply decision by a non-industrial private forest owner on a timber market where all timber is sold as stumpage. Stumpage prices (p_t) are assumed to follow a stationary first-order auto-regressive process over time periods, denoted by *t*:

$$\log(p_{t+1}) = \alpha + \log(p\quad) + u_t; \qquad 0 \leq \beta \leq 1 \tag{1}$$

$$E(u_t \mid I_t) = 0 \tag{2}$$

where α and β are constant parameters and the stochastic term u_t is assumed i.i.d. (I_t denotes the information set at time *t*).

Further assume that the relative growth of the value of a timber stand at constant prices is a positive decreasing function $g(b_t)$ in the age of the stand (b_t). The forest owner's opportunity cost of capital is r_t. The forest owner can choose between selling the stand now (time *t*) or later (time *t*+1). Given the information available at time *t*, she is indifferent if

$$g(b_t) = (1-\beta)\log(p_t) + r_t - \alpha \tag{3}$$

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[2] However, the price volatility affects the harvesting threshold. A higher volatility increases the rotation age (see also Amacher et al. 2009, Ch. 11).

[3] This statement should be modified in several ways. First, the analysis of Clarke and Reed (1989) of tree cutting under uncertainty should be compared to the deterministic setup of Knut Wiksell in his famous wine-aging problem since (unlike in the 1849 study by Martin Faustmann) only one generation of trees is considered. For a more complete treatment, see Amacher et al. (2009, Ch. 11). Second, David and Cairns (2012) show that the theory of stopping under uncertainty is a generalization of the rule under certainty, and that both can be united as special cases of an "r-percent stopping rule".

where b_t is the minimum age of a stand that will be sold at time *t*. Differentiating this condition with respect to p_t and b_t gives

$$\frac{db}{dp_t} = \frac{(1-\beta)}{g'\,p_t} \leq 0 \tag{4}$$

The short run supply is increasing when the minimum cutting age is lowered. From (3) and (4) we see that short run supply is increasing in p_t and *r* (since the marginal age is lowered). Notice that supply would be perfectly inelastic if $\beta = 1$, thus when $\log(p_t)$ follows a random walk with drift.

3. Stumpage prices in Sweden 1909 – 2011

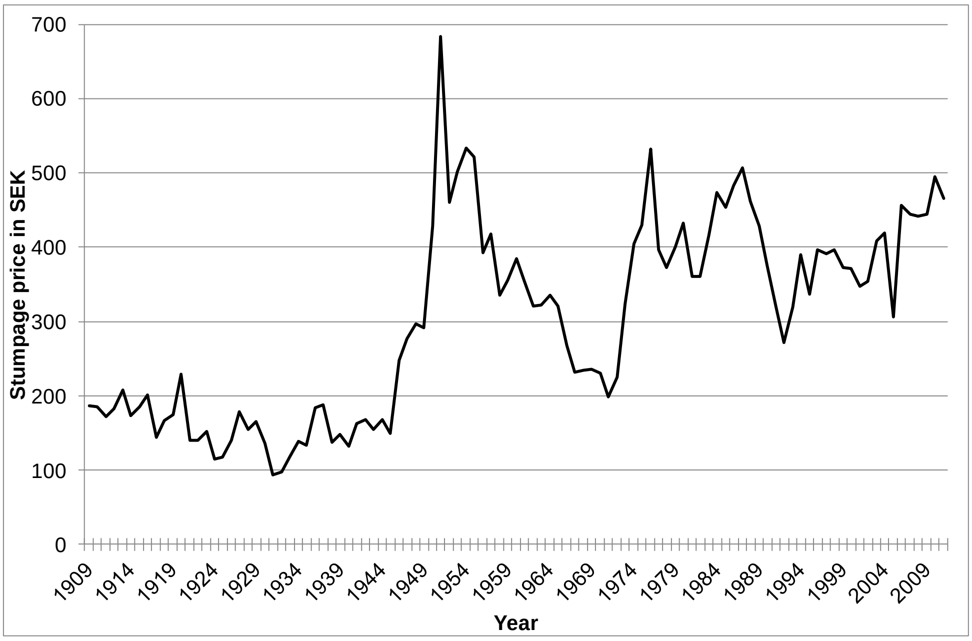
The modern market for stumpage emerged at the turn from the 19th to the 20th century from legislation intended to reduce emigration from rural areas[4] by protection of the non-industrial private ownership to forest lands. The maximum duration of harvesting rights was limited to 20 years in 1889 and further down to five years in 1903, and from 1905 forest companies were prevented from buying forest land in the northern part of the country (from 1925 in the whole country). This established a basic structure for the stumpage market that still remains, with a quarter of a million non-industrial private forest-owners on the supply side and a five years limit for the right to exercise the harvesting rights that are transferred through stumpage trade. Markings and assessments of timber volumes are as a rule made by an independent third party, and the stumpage is sold in a simple auction, often organized as a sealed-bid one-shot auction.

In Figure 1 we show real annual stumpage prices in Sweden from 1909 – 2011. Prices for the period 1909-1955 were collected by Streyffert (1960); for the period 1956 – 2002 by the

[4] Sweden lost 1.2 million emigrants, mostly to the U.S., which can be compared to the total population year 1900 of 5.2 million (Stråth 2012).

Swedish National Forest Board (Annual Yearbook of Forestry, various issues); and for 2003 – 2011 by the stumpage broker Rotpostmäklarna AB (2012). Prices are deflated by the Consumer Price Index (CPI).[5] The data are also listed in the Appendix.

Figure 1 Annual real stumpage prices in Sweden 1909 – 2011, SEK/m3 (standing volume)



Note: The price series is deflated by the consumer price index with 2011 as the base year.

Source: Streyffert (1960), Yearbook of Forest Statistics (various issues) and Rotpostmäklarna AB (2012).

The prices shown in Figure 1 are to large extent genuine annual prices because of the marked seasonality of logging operations.[6] Contracts were traditionally signed in the autumn, felling made wintertime and river floating in the spring. Nowadays, logging machines can be used throughout the year, but stumpage trade still peaks in the autumn (Bjerke, 2012). Seasonal conditions, such as spring thaw and the short vegetation period, remain important considerations in forestry planning.

[5] Before 1935 prices are deflated by the so called Living Cost Index computed by the National Social Welfare Board.

[6] Prestemon et al. (2004) demonstrate, using U.S. southern states stumpage price quarterly data from 1977 to 2002, that temporal averaging in reported timber prices can reduce the statistical power of unit-root tests. Quarterly stumpage price data does not exist for Sweden.

Parts of this price series has been previously analyzed by Streyffert (1960) and Hultkrantz (1995). Streyffert had access to data up to 1955. The series he observed was thus dominated by the hike from a minimum in 1931 during the Great Depression to the 1951 peak during the Korea war. Hultkrantz (1995) studied the data up to 1990. He applied the Perron (1989) unit root test allowing for a structural (level) break at the end of WW2 and concluded that stumpage prices are stationary.

A reason for expecting a structural break at the end of WW2 is that the peace released a series of structural changes of the rural economy of the forest regions that had been held back by the Great Depression and the war. One of these was the unification of the national labor market, which previously was a dual market characterized by low productivity and excess supply in the remote rural forest regions. Most non-industrial forest owners had subsisted on agriculture on small parcels, combined with seasonal work in forestry, timber-floating, sawmills, etc. As rural labor was relatively cheap, forestry relied until the end of the war on muscle power (by men and horses), which was later than in, for instance, Canada. After WW2 differences between rural and urban labor markets rapidly leveled out, through commuting and intensified urbanization. Small farms were merged into larger production units and many non-industrial private forest owners got regular employment. Forest owners therefore became less dependent on a continuous stream of forest revenue.

Also, a majority of the non-industrial forest owners joined the co-operative forest owners´ associations for the wartime distribution of fuel wood. Therefore, the large forest companies that had organized buyers´ cartels for delivery sales of pulpwood and saw logs, based on regional separation of the timber markets in the 1920’s, from the end of the war had to negotiate price lists for delivery sales with much stronger representatives of the forest owners than before the war.

On the markets for delivery sales of pulpwood and saw timber, round wood prices were announced annually after negotiations between the buyers´ and sellers´ organizations until the mid- 1990´s when a new Competition Act made horizontal price co-operation on both sides of the market illegal. Stumpage markets, however, remained an “un-organized” competitive fringe before and after WW2, on which forest owners sold timber individually, outside of the co-operative structures, and where large forest companies had to compete for timber with small and medium size sawmills that were active on local markets only; see Bergfors et al. (1989).

Visual inspection of Figure 1 suggests that the development after WW2 can be divided into three separate stages: 1945 – 1951; 1951-1972; and 1972 – 2011. The first period is the first after-war years marked by the loss of production capacity of forest industry in other European countries, ending with the Korea war inflation episode with rocketing prices on many raw materials, including timber. During the second period the Swedish forest industry met more fierce completion from restored and modernized production capacity in other countries. However, this spurred a wave of structural changes of the Swedish industry and a technological shift (much driven by domestic innovation) towards the use of timber harvesting machines, leaving “no feet on the ground”. Later, these machines were amended by computer and sensor technologies, etc.

Clearly, during the more than hundred years spanned by our data several technological and institutional changes have occurred. When testing for unit root in this data series the possibility of breaks therefore needs special attention. In next section we will have a new look on testing for non-stationarity in the stumpage-price series account for a break at the end of WW2, while we in the subsequent section allow for an endogenously determined break point.

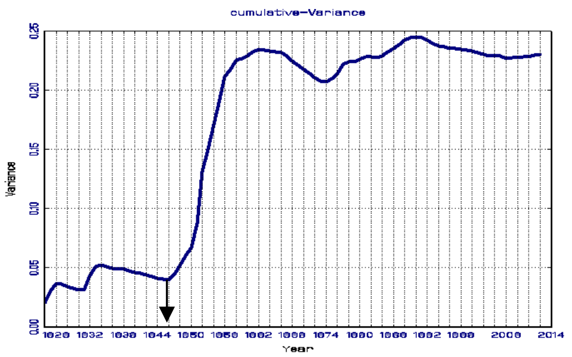
4. Testing for unit root

The classical unit roots tests (Dickey-Fuller test and Augmented Dickey-Fuller test) tend to *not reject* the unit root of a time series with changing mean or breaking trend, that is, they have low power in the case of a break. Perron (1989) considered three versions of hypothesis testing for unit roots and structural change. However he used as test model the case of *random walk with drift and trend*, and also with an exogenously determined break point. The approach by Perron (1989) of dating the structural break was criticized and Christiano (1992) argued that by exogenously dating the break points the researcher will build in problems of e.g., sample selection. The following literature such as Zivot and Andrews (1992), Perron and Vogelsang (1992), Perron (1997), Lumsdaine and Papell (1997) and Bai and Perron (2003) has therefore used various methods to endogenously find the breakpoint(s) in the data series.

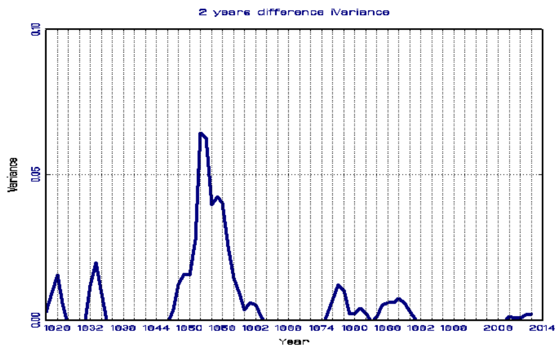
According to Figure 2 the Swedish stumpage price series does not show evidence of following a model with *linear trend in the mean*. However, the sample autocorrelation function (SACF) in Figure A1 shows that there is persistence in data, which may indicate either that we should model the price series with a *linear trend in the variance* or with a structural break around 1945. To be able to decide which model to use, we need to analyze the variance of the stumpage price series further.

Figure 2 Cumulative and difference in variance of annual stumpage prices (in logs) in Sweden 1909-2011 to detect structural breaks

Panel a) Cumulative variance



Panel b) 2 years difference in variance



Note: Panel a) refers to cumulative block variance using and panel b) refers to the difference in variance between two years.

Assume the data generating process starts in 1909 (the first year of observation) and that we allow 15 years before we start analyzing data (1924). We use this dataset to calculate the sample variance, s_1^2. The following year we repeat our analysis by taking the sample mean, s_2^2, with 16 observations (1909-1925), and continue repeating the analysis up until 2011. The cumulative variance with this procedure is shown in Figure 2 panel a). From this figure a clear pattern evolves where the variance clearly starts to increase in 1946. However, since this increase in the variance stops after 1960 and the variance stabilizes at a higher level than pre-1946, this is rather evidence of a structural break in the time series than evidence of a linear trend in the variance. Thus, it appears as if the stumpage price series should be modeled with a structural break in the non-zero mean at the year 1946. For robustness we also calculate the difference in variance between two subsequent years, which is reported in Figure 2 panel b). Once again, data suggest that there is a significant difference in the mean starting in 1946 indicating a structural break. This increase in the mean is more than 5 percent during that time, which can be used as a rule of thumb.

Let T_B be the break time, a breakpoint that is given a by the above analysis of the variance in the price series, of the sample period T. In our data T_B = 1946, which is the 38th observation out of a total of 103 annual observations. However, since our model is different from the setup in Perron (1989) we solve this problem by simulating critical values by a Monte Carlo experiment.

Consider that the time series y_t is a random walk:

$$y_t = y_{t-1} + e_t \text{, where } e_t \sim i.i.d(0,\sigma^2) \tag{1}$$

Equivalently, we write (1) as:

$$y_t = y_0 + \sum_{t=1}^{T} e_t \tag{2}$$

Assuming now that the series is generated by the following process:

$$y_t = \left[y_0 + \sum_{t=1}^{T_B} e_t \right] || \left[(y_0 + c) + \sum_{t=T_B+1}^{T} e_t \right], \tag{3}$$

where c is an integer (positive in our case) and that $e_t \sim i.i.d(0,\sigma^2)$. The merge of the full time series gives y_0 up until the T_B point, where the time series restarts with $y_0 + c$. However, (3) can equivalently be written as:

$$y_t = \left[(y_0 + cD) + \sum_{t=1}^{T} e_t \right], \tag{4}$$

where D is a dummy variable with T_B number of zeroes and $(T - T_B)$ ones.

Unit root test procedure

The early studies using endogenous dating of the break did not allow for the break in the null hypothesis, which means that those tests may be biased towards suggesting evidence of stationarity with breaks (Lee and Strazicich, 2003). Note that Perron (1989) allows for a break both under the null and alternative hypotheses. It is nowadays well known (Hamilton, 1994) that when testing for unit roots it is important to specify the null and alternative hypotheses appropriately based on the type of the data at hand. So based on our data we use a constant term plus a dummy variable (for the break) in the regression. The test regression is given by

$$y_t = c + \phi y_{t-1} + dD + \varepsilon_t \tag{5}$$

and the hypotheses to be tested are

$$H_0 : \phi = 1 \qquad \Rightarrow \qquad y_t = I(1) \text{ without a drift}$$

$$H_A : |\phi| < 1 \qquad \Rightarrow \qquad y_t = I(0) \text{ with non-zero mean.}$$

The test statistic we use is the usual:

$$t_\phi = \frac{\hat{\phi} - 1}{SE(\hat{\phi})} \tag{6}$$

To test the unit root hypothesis we use Monte Carlo simulated variables. We generate 103 observations based on model (4) with start value y_0 given by the first value of our series and c for a dummy variable equal to $(y_{38} - y_1)$. Then OLS is used to estimate the test regression (5) and the test statistic t_ϕ (6). Repeating these steps a large number of times, N_{MC} (in our Monte Carlo simulations, N = 10000), we are able to approximate the finite-sample distribution of the unit root test.

By taking the (1–α) quintile of the approximate finite-sample distribution of t_ϕ, we obtain the α-level "critical values" ($c_{t\alpha}$) given below in Table 1.

Table 1 Monte Carlo simulated critical values

Level of significance	**Critical value**
1 percent	-3.9782535
5 percent	-3.3549926
10 percent	-3.0484206

Table 2 Test results with dependent variable is the first difference of stumpage prices

	Coefficient	**Std. Error**	**t-ratio**
constant	1.512	0.367	4.12
break-dummy	0.262	0.072	3.64
y_{t-1}	-0.298	0.072	-4.11
Adj. R2	0.128		
Durbin Watson	1.997		
Akaike	-72.526		
No. of obs. (*T*)	102		

Note: $y_t = c + \phi y_{t-1} + dD + \varepsilon_t$ is equivalent to estimating $\Delta y_t = c + (\phi - 1) y_{t-1} + dD + \varepsilon_t$.

According to the results presented in Table 2, the observed t-value is -4.11, which is less than the simulated critical value of -3.978. This means that we can reject the null hypothesis of non-stationarity at the 1-percent significance level. i.e., the Swedish annual stumpage price series does not show evidence of a unit root.

4.1 Endogenously given break point

To test the robustness of our results of unit root testing of the Swedish stumpage prices we will apply two readily available unit root tests allowing for a break, namely the Zivot and Andrews (1992) approach and the approach by Clemente et al. (1998). The latter test builds on Perron and Vogelsang (1992). Since Figure 2 shows that our data series does not appear to have a trend element the relevant model is to test for a break in the mean.

Table 3 Unit root test with one endogenously dated break point in the mean

Test	**Break point**	**Minimum statistics**	**Critical value (5-percent level)**
Zivot and Andrews (1992)	1946	-4.977	-4.80
Clemente et al. (1998)	1944	-5.399	-4.27

Note: The Clemente et al. (1998) test refers to the innovational outlier (IO) model with a break in the mean.

According to both test results presented in Table 3, the Swedish stumpage price series is a stationary process with a break. Using the Zivot and Andrews (1992) approach gives us the same break point as suggested by Figure 2, while the Clemente et al. (1998) test suggests an earlier break point of 1944.

6. Discussion

Our results thus suggest that the long term development of Swedish stumpage prices is stationary mean reverting, when account is taken of a break at the end of WW2. When the break is modeled as a step (intercept) break it comes 1946, if instead modeled as an innovational outlier (which means that the full effect builds up over some time) it seems to start in 1944.

Very long time series for stumpage prices are rare. However, in a conference paper Lutz (1999) presents a time-series for U.S. southern pine stumpage prices from 1890-1996 that was provided by the Hancock Timber Resource Group. From ADF and Phillips-Perron tests he concludes that the series over the whole range is “probably not stationary”. He makes this claim deliberately vague as he observes that there are several sharp spikes in the data that make the results from these tests uncertain as they do not consider possible break points. In fact, inspection of a graph of the time series presented in his paper (Figure 5) suggests a rather close resemblance with the Swedish data (from 1909 to 1996), i.e., there is no clear trend before or after WW2, but a strong upward shift in the first peace time years.

Four series of quarterly stumpage price data 1973 – 1997 for different tree species of the U.S. Pacific Northwest national forests were analyzed by Saphores et al. (2000). They test for unit root with a Perron test accounting for a structural break because of a statistical definition change in 1984. This test does not reject non-stationarity in any of the four series. Similar results were held by Prestemon et al. (2004) in analysis of quarterly data of stumpage prices for U.S. southern pine sawtimber, southern pine pulpwood and hardwood pulpwood 1977 – 2002. Both these studies analyze data during a much shorter time span than our data. In fact, ADF test for subsets of our data over similar time periods reject or do not reject unit root, depending on exactly what start year is used. This indicates that such results are not very robust.

Another study that used a long time series is Bayazidi and Yoshimoto (2011). They study Finnish data from 1900 to 2007 for the average yearly price of softwood logs (i.e., a “gross” timber price). Interestingly, they find that ADF and Phillips-Perron tests reject unit root, while the Kwiatkowski-Phillips-Schmidt-Shin test does not reject trend stationarity. This is in contrast to results from tests of series of monthly data from 1988 – 2008 that indicate unit root. This study, also using a different kind of data, thus points in a similar direction as our study.

7. Conclusions

As reviewed in the textbook on forest economics by Amacher et al. (2009), much recent work on the theory of efficient forest management under uncertainty is based on the assumption that the logarithm of stumpage price follows a random walk (geometric Brownian motion). This assumption is supported in some empirical analysis, although some other studies have found stumpage prices to exhibit mean reversion.

However, the conventional tests for unit root used in a large part of the empirical research can be misleading for two reasons. First, their statistical power is low, so a large sample of observations is needed. For this reason, the tests are not very robust to changes of the precise delimitation of the data set, in particular, the choice of start year may be pivotal. Second, structural breaks can disturb the tests results.

In this work, we use a consistent data set spanning stumpage prices over more than hundred years. We find that (real-valued) stumpage prices have no time trend and are mean reverting, thus we reject the random walk conjecture. A structural break occurred, though, at the end of WW2. The behavior of the cumulative variance of the data that was displayed in Figure 2 indicates that both these results are robust.

For future research, it would be very interesting to see whether our findings can be corroborated by similar analysis of long time-series data for stumpage prices at other places; for instance such data seems to be available for southern U.S. pine stumpage sales. For much broader research, our results raise questions on how to understand the long-term development of stumpage prices. What are the forces that seem to preserve long-term price stability on a market that over a century (except for some time during WW1 and WW2) has been fully exposed to the world market for forest products and thereby to numerous technological shocks, business-cycle fluctuations, and global structural change?

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Appendix

Table A1 Swedish annual stumpage prices in current and fixed 2011-values (SEK/m3. standing volume) 1909-2011

Year	Current price	Fixed price	Year	Current price	Fixed price	Year	Current price	Fixed price
1909	3.84	186	1944	8.63	168	1979	121.00	401
1910	3.86	185	1945	7.67	149	1980	147.00	432
1911	3.69	172	1946	12.94	248	1981	135.00	360
1912	3.95	182	1947	15.09	277	1982	147.00	360
1913	4.54	208	1948	16.70	297	1983	184.00	416
1914	4.09	173	1949	16.51	291	1984	226.00	474
1915	4.99	185	1950	26.50	428	1985	229.00	454
1916	6.48	201	1951	47.30	683	1986	254.00	483
1917	6.41	144	1952	33.07	460	1987	280.00	507
1918	9.33	167	1953	36.26	501	1988	270.00	461
1919	10.25	174	1954	39.18	533	1989	272.00	428
1920	12.37	229	1955	39.76	521	1990	262.00	375
1921	6.25	140	1956	31.44	393	1991	236.00	320
1922	5.49	140	1957	34.93	418	1992	207.00	271
1923	5.82	152	1958	28.72	335	1993	252.00	319
1924	4.44	115	1959	31.12	355	1994	315.00	390
1925	4.48	117	1960	34.70	384	1995	276.00	337
1926	5.23	140	1961	33.14	354	1996	326.00	396
1927	6.68	179	1962	31.17	321	1997	323.00	391
1928	5.75	154	1963	32.23	322	1998	328.00	397
1929	6.00	165	1964	34.96	335	1999	310.00	372
1930	4.78	136	1965	35.49	321	2000	314.00	371
1931	3.22	94	1966	31.05	267	2001	301.00	347
1932	3.31	98	1967	27.75	231	2002	313.00	354
1933	3.91	117	1968	28.78	234	2003	364.64	408
1934	4.69	139	1969	30.18	235	2004	376.41	419
1935	4.55	133	1970	31.69	230	2005		306
1936	6.44	184	1971	29.31	199	2006	420.90	456
1937	6.74	188	1972	35.23	225	2007	421.36	444
1938	5.06	137	1973	54.89	324	2008	425.59	442
1939	5.89	148	1974	75.16	404	2009	430.34	444
1940	5.99	132	1975	88.00	429	2010	488.26	495
1941	8.06	162	1976	121.00	532	2011	465.91	466
1942	8.68	168	1977	100.00	397			
1943	7.92	154	1978	102.00	373			

Note: Due to the vast and extraordinary effects of the storm Gudrun in 2005 no stumpage price is reported by Rotpostmäklarna AB. Prices 1909-1934 are deflated by the so called Living Cost Index computed by the National Social Welfare Board and prices 1935-2011 are deflated by the Consumer Price Index computed by Statistics Sweden.

Source: Streyffert (1960): 1909-1955; Swedish National Forest Board (Annual Yearbook of Forestry. various issues): 1956-2002; Rotpostmäklarna AB: 2003-2011.

Figure A1 Sample autocorrelation and partial autocorrelation functions of annual Swedish stumpage prices (in logs)

Figure A1a) Annual stumpage prices in logs

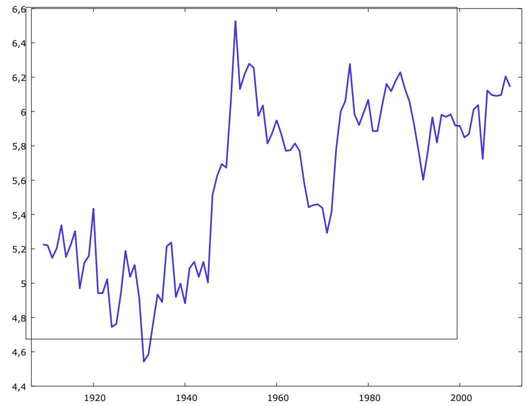

Figure A1b) SACF and SPACF