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A behavioral timber supply function is derived assuming a representative nonindustrial private forest owner who decides an optimal date for even-age harvesting simultaneously with his optimal life cycle consumption and savings. Capital markets are perfect and forests are valued *in situ*. The effects of intertemporal changes in the exogenous variables on the annual harvest are estimated using a fixed-effects tobit model and a sample of 119 Finnish nonindustrial private forest owners in 1982-1991. The incidental parameter bias is demonstrated to be small in Monte Carlo simulations. An intertemporal effect of stumpage price is positive while the negative intertemporal effect of nonforestry income may indicate capital market imperfections. The effects on the harvest of permanent parametric shocks in explanatory variables are studied by regressing the estimated individual effects on the study period averages of the explanatory variables. The parametric effect of the stumpage price is negative. Nonforestry income has a positive parametric wealth effect, supporting the hypothesis that forests are valued *in situ*, but it is smaller in absolute value than the negative intertemporal effect.

Paneeliaineistoon perustuva empiirinen puuntarjontatutkimus ei tyypillisesti erottele toisistaan poikkileikkaus- ja aikasarjainformaatiota. Puun tarjonnan riippuvuutta metsänomistajakohtaisista tekijöistä on selitetty kahden periodin rakenteeseen perustuvalla kulutus/säästämismallilla, johon on lisätty metsän ikäluokkia erottelematon uusiutuvien luonnonvarojen käyttöä kuvaava osamalli. Kuitenkin jo Faustmannin optimirotaatiomallista johdettavien hypoteesien mukaan puun hinnan parametrinen nousu alentaa puun tarjontaa, kun taas intertemporaalinen nousu aiheuttaa lyhyellä aikavälillä tarjonnan kasvun. Tämän seurauksena puun tarjontaa koskevia hypoteeseja ei voi perustellusti testata tekemättä empiiristen havaintojen kohdalla eroa poikkileikkaus- ja aikasarjainformaation välillä. Vastaavia ongelmia aiheutuu tutkittaessa esimerkiksi varallisuuden ja puun tarjonnan yhteyttä. Lisäksi tutkittaessa puun tarjontaa suomalaisella aineistolla on perusteltua soveltaa teoreettista kehikkoa, joka kuvaa puuntarjontapäätöksen riippuvuutta kaadettavan puuston iästä.

Tässä tutkimuksessa sovelletaan metsänomistajakohtaiset tekijät sisältävää optimirotaatiomallia. Metsänomistajan oletetaan maksimoivan elinaikaista hyötyä, joka riippuu kulutusmahdollisuuksista ja metsän ei-puuntuotannollisesta arvosta. Pääomamarkkinoiden oletetaan olevan täydelliset. Perinnönjättömotiivi ilmenee ainoastaan rahavarallisuuden ei-negatiivisuusehtona elinkaaren lopussa. Puun hinnan oletetaan vaihtelevan esimerkiksi suhdannevaihteluiden seurauksena. Mallin hypoteesien mukaan puun hinnan intertemporaalinen nousu lisää puun tarjontaa, eli rationaaliset metsänomistajat pyrkivät ajoittamaan puunmyynnin ajankohtiin, joissa puun hinta on korkeimmillaan. Puun hinnan parametrisen kasvun (esim. alueelta toiselle siirryttäessä) voidaan kuitenkin ennustaa alentavan puun tarjontaa. Vastaavasti täydellisten pääomamarkkinoiden vallitessa ei metsätalouden ulkopuolisten tulojen ajallisella vaihtelulla pitäisi olla vaikutusta puun myynnin ajoitukseen. Sitävastoin tulojen parametrisen kasvun pitäisi lisätä puun tarjontaa.

Tutkimuksen aineisto käsitti 119 eteläsuomalaista maanviljelijämetsänomistajaa, joita oli haastateltu vuosina 1985, 1986 ja 1991. Aineisto sisälsi tiedot vuosittaisesta puunmyynnistä, puuvarannon kehityksestä ja esim. metsätalouden ulkopuolisista tuloista ja varallisuudesta.

Johdetun puuntarjontamallin intertemporaaliset vaikutukset estimointiin käyttäen kiinteiden vaikutusten tobit mallia. Puun myyntien vaihtelua metsänomistajien välillä eli selittävien muuttujien parametristen muutosten vaikutuksia puun tarjontaan puolestaan tarkasteltiin käyttämällä tobit mallilla estimoituja metsänomistajakohtaisia vakiokertoimia, joita selitettiin eksogeenisten muuttujien tutkimusperiodille lasketuilla metsänomistajakohtaisilla keskiarvoilla. Tulosten luotettavuutta tutkittiin empiiristä aineistoa vastaavan keinotekoisen aineiston avulla.

Tulokset osoittavat, että paneeliaineiston käyttö erottelematta muuttujien intertemporaalisia ja parametrisia vaihteluita johtaa teoreettisten hypoteesien hylkäämiseen. Sitävastoin puun hinnan intertemporaalisen vaihtelun ja puun tarjonnan tarkastelu osoittaa, että metsänomistajat lisäävät puun tarjontaa hintahuipuissa juuri kuten teoreettinen malli ennustaa. Lisäksi puun tarjonta näyttää riippuvan käänteisesti metsätalouden ulkopuolisten tulojen vaihteluista. Tämä on vastoin täydellisiin pääomamarkkinoihin perustuvan deterministisen mallin ennusteita, mutta on tulkittavissa rationaaliseksi tarjontakäyttäytymiseksi esimerkiksi luotonsäännöstelyn oloissa. Tarkasteltaessa hinnan parametrisiä muutoksia saatiin tulos, jonka mukaan puun tarjonta on kääntäen verrannollinen puun hintatasoon. Tämä tulos tukee optimikiertoaikamallin hypoteesia, jonka mukaan puun pitkän aikavälin tarjontakäyrä on laskeva. Metsätalouden ulkopuolisten tulojen parametrinen kasvu lisää puun tarjontaa. Tämä tukee teoreettista hypoteesia, jonka mukaan metsänomistajan varakkuus implikoi pitempää rotaatioperiodia ja suurempaa vuotuista puun tarjontaa. Tulojen negatiivinen intertemporaalinen vaikutus on kuitenkin varallisuusvaikutusta voimakkaampi.

Key words: timber supply, forest rotation models, panel data, censored dependent variables.

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## Content



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# 1 Introduction

Several empirical studies have demonstrated that nonindustrial private forest (NIPF) owners' harvesting decisions depend on owner characteristics that should not affect the supply if forest owners were maximizing the present value of harvest income (Binkley 1981, Carlén 1990, Dennis 1988 and 1990, Hyberg and Holthausen 1989, Kuuluvainen and Salo 1991, Kuuluvainen et al., 1996). Many of these studies use a Fisherian consumption-savings model augmented with renewable stock dynamics as a theoretical framework to model the "short-term" timber supply (e.g. Aronsson 1990). However, Hyberg and Holthausen (1989) show that forest owner characteristics can also be incorporated into the Faustman model.

The empirical analyses have used cross-section (e.g. Aronsson 1990, Carlén 1990, Dennis 1988) or pooled cross-section/time-series data (e.g. Hyberg and Holthausen 1989, Dennis 1990, Kuuluvainen et al. 1996). In pooled estimation, cross-section and intertemporal effects are added together. Therefore, when pooled data have been used it has not been clear whether the parameter estimates are the effects of cross-sectional variation in the exogenous variables or are due to the intertemporal variation. In the former case the effects should be interpreted as the effects of (unexpected) parametric shocks in exogenous variables, whereas in the latter case the effects describe (expected) intertemporal development over e.g. a business cycle.

The contribution of this paper is that we derive hypotheses concerning both the parametric and intertemporal effects using a utility-based rotation model developed by Tahvonen (1997) and test them separately in estimating the timber supply equation. For example, according to the basic Faustmann model cross-sectional variation in timber prices has a negative effect on harvest due to a shorter rotation period. The short-run effect of an unexpected price increase is positive, as it makes some stands overmature. However, in our data sample, more relevant price movements affecting intertemporal short-run variation in timber harvest are the anticipated price fluctuations over e.g. a business cycle. These are seldom studied in optimal rotation literature. We show that in a utility-based Faustmann-type optimal rotation model high timber prices are associated with high observed harvest. Thus, the sign of the intertemporal effect is positive as is the sign of the short-run effect of an unanticipated price change.

In order to differentiate the intertemporal and parametric effects empirically, a two-step estimation procedure similar to Heckman and MaCurdy (1980) (see also MaCurdy 1981) is used. The first stage provides estimates for the effects of intertemporal variation in the exogenous variables on the volume harvested using fixed-effects models. The effects of permanent or temporary parametric shocks in timber prices, exogenous nonforestry income, and owner characteristics on harvest is studied using individual constant terms estimated in the first stage. The incidental parameter bias may be a problem due to the small number of time periods in the sample and the dependence of the commom slopes on incidental parameters in a maximum likelihood tobit model. Using Monte Carlo simulation, we demonstrate that the problem may not be serious in the present sample (see also Heckman 1981, Cassel et al. 1996).

According to the empirical results, the intertemporal effect of timber price is positive and that of exogenous nonforestry income is negative. The latter may indicate the effect of an imperfect capital market. Parametric effects of these variables take signs opposite opposite to those of the intertemporal effects as predicted by the rotation model with *in situ* benefits. Therefore, the long-run timber supply function in the present sample of 119 Finnish NIPF owners is in fact backward bending, as predicted by the basic Faustmann model. The short-run timber supply function has the normal positive slope.

Section 2 the presents the theoretical model and the estimable equation for timber supply. Section 3 describes the institutional environment and the data. The empirical results are given in Section 4, and Section 5 concludes the paper.

# 2 Even-age harvesting and life cycle consumption-savings decisions

## 2.1 Theoretical model

Theoretical timber supply models for nonindustrial private forest owners are normally either Faustmann-type profit maximization models without consumption decisions or two-period lumped parameter renewable resource models with consumption-saving dynamics. Because Finnish forestry is based on even aged management, we prefer a Faustmann-type rotation model, but we also want to include forest owner-specific characteristics and consumption-savings decisionmaking. In the following we apply the *in situ* version of such a model developed by Tahvonen (1997).

In boreal forests the rotation period varies from 50 to 130 years. By contrast, an average Finnish NIPF owner holds his forest land only about 30 years between acquisition it from his parents and leaving it to his heirs. We neglect the market for forest land because of its lack of relevance in our study sample. The forest owner's problem is to choose an optimal year for cutting the forest within his lifetime while simultaneously making his consumption-savings decisions, acting in capital markets and receiving nonforest income. For simplicity, our theoretical model assumes that each forest owner has only one even-age stand. We first formulate his optimal comsumption-savings decisions for the period after the stand is cut. The solution for this problem yields a value function which we can use for the problem before the stand is cut to determine the optimal cutting moment and timber supply.

Let $U(c)$ denote a strictly concave utility function with inada conditions, where c is consumption. A(x) denotes the value of the stand *in situ*. Stand volume is given by x and $A' > 0$, $A'' \geq 0$. The forest growth function $F(x)$ is concave and $F(0) = F(\overline{x}) = 0$, $\overline{x} > 0$. Denote the date when the stand is cut by $t_1$ and the length of the decisionmaker's life cycle as a forest owner by $T$. The annual interest rate is $\rho$, the subjective time preference $\delta$, nonforest income m, stumpage price p, planting costs w, and the level of nonforest assets a. At the cutting moment there is a jump down in the stand volume and a jump up in nonforest assets. We use the notation $t_1 = t_1^-$ when $t_1$ approaches $t_1$ from below.

The problem for the period after the stand is cut is

$$W_1 = \max_{\{c \geq 0\}} \int_{t_1}^{T} [U(c) + A(x)] e^{-\delta t} dt, \tag{1}$$

$$\text{s.t. } \dot{a} = \rho a - c + m, \quad a(t_1) = a(t_1^-) + px(t_1^-) - w, \tag{2}$$

$$a(T) \geq 0, \tag{3}$$

$$\dot{x} = F(x), \ x(t_1) = x_{t_1}, \tag{4}$$

where $x_{t_1}$ is the volume of the stand just after planting. In equation (1) we have neglected possible bequest motives for forest and nonforest assets[1]. However, we have assumed in (3) that the owner is not willing to leave a debt to his heirs. Equation (2) shows the value of nonforest assets after the stand is cut. This value equals the level of assets before cutting, $a(t_1^-)$, plus the net income from the cut, $px(t_1^-) - w$. Planting costs w are subtracted because for our sample replanting is required by law. The current value Hamiltonian and necessary conditions for an optimum are [2]:

$$H = U(c) + A(x) + \lambda(\rho a - c + m) + \varphi F(x), \tag{5}$$

$$U'(c) - \lambda = 0, \tag{6}$$

$$\dot{\lambda} = \lambda(\delta - \rho), \tag{7}$$

$$\dot{\varphi} = -A'(x) + \varphi[\delta - F'(x)], \tag{8}$$

$$\varphi(T) = 0, \tag{9}$$

together with conditions (2)-(4). The problem (1)-(4) satisfies the requirements of theorem 9 in Seierstad and Sydsaeter (1987, p. 213), implying that the optimal value function is differentiable w.r.t. initial assets and length of time horizon. Denote the optimal value function

[1] An infinite horizon version of this model is studied in Tahvonen and Salo (1997).

[2] Seierstad and Sydsaeter (1987, p. 182, theorem 3). By theorem 4, p. 182, the conditions are sufficient.

by $W_1[a(t_1^-), x(t_1^-), t_1]$. The problem for the period before the stand is cut is

$$\max W_2 = \int_0^{t_1} U(c)e^{-\delta t} + W_1\left[a(t_1^-), x(t_1^-), t_1\right] \tag{10}$$

$$\text{s.t.} \qquad \dot{a} = \rho a - c + m, \qquad a(0) = a_0 \tag{11}$$

$$\dot{x} = F(x), \; x(0) = x_0, \tag{12}$$

where $a_0$ and $x_0$ are the initial levels of nonforest assets and the volume of the stand. The Hamiltonian is again given by (5) and the necessary conditions include (6)-(8), (11), (12) and

$$\lambda(t_1^-) = e^{\delta t_1} \partial W_2 / \partial a(t_1^-), \tag{13}$$

$$\varphi(t_1^-) = e^{\delta t_1} \partial W_2 / \partial x(t_1^-), \tag{14}$$

$$H(t_1^-) = H(t_1) - \lambda(t_1)\dot{p}x(t_1^-). \tag{15}$$

Note that the RHS of (15) includes two different effects[3] that the choice of $t_1$ has on the value of $W_1$. The longer the growing period, the shorter the period after the cut, as reflected by $H(t_1)$. However, if timber price is not constant, an increase in $t_1$ changes the value of the cut and the initial level of nonforest assets after the cut, as reflected by $-\lambda(t_1)\dot{p}x(t_1^-)$.

Applying theorem 9 in Seierstad and Sydsaeter (1987, p. 213), we obtain by equation (13) that $\lambda(t_1^-) = \lambda(t_1)$, and by equation (14) that $\varphi(t_1^-) = \lambda(t_1)p$. Continuous $\lambda$ (and c) together with (5) and (15) now leads to

$$\begin{aligned} &p(t_1^-)\lambda(t_1)\left\{F\left[x(t_1^-)\right] - x(t_1^-)\left[\rho - \dot{p}(t_1^-)/p(t_1^-)\right] + \rho w / p(t_1^-)\right\} \\ &+ A\left[x(t_1^-)\right] - A(x_{t_1}) - \varphi(t_1)F(x_{t_1}) = 0. \end{aligned} \tag{16}$$

It is possible to show that $A\left[x(t_1^-)\right] - A(t_1) - \varphi(t_1)F(x_{t_1}^-) > 0$, implying that *in situ* benefits lengthen the rotation (Tahvonen 1997). The terms $p(t_1^-)\lambda(t_1)F\left[x(t_1^-)\right] + A\left[x(t_1^-)\right]$ on the LHS of (16) denote the value of letting the initial stand grow marginally longer. The term

[3] Seierstad&Sydsaeter (1987) theorem 4 and note 2, pages 182 and theorem 9 page 213.

$p(t_1^-)\lambda(t_1)\left\{-x(t_1^-)\left[\rho-\dot{p}(t_1^-)/p(t_1^-)\right]+\rho w/p(t_1^-)\right\}$ denotes the interest cost of receiving the net gain from the cut later in time, and $-A(x_{t_1})-\varphi(t_1)F(x_{t_1})$ is the loss in second period *in situ* utility due to a marginal increase in $t_1$.

Due to a varying timber price and *in situ* benefits, it is impossible to rule out multiple solutions for equation (16). In these cases the globally optimal solution is the cutting moment that gives the highest value for $W_1$. By using this fact, we can compute the comparative static properties for the optimal rotation periods and thus show how timber supply depends on various forest owner-specific and other factors. Note that the coefficient for $\lambda(t_1)$ in (16) is negative, implying that for any local maximum the rotation period is shorter, the higher the marginal value of consumption. This is intuitive because higher marginal value of consumption makes it more costly to deviate from the cutting moment that would maximize the pecuniary income from harvesting the stand. To obtain the comparative static results in more detail, it is possible to solve the marginal value of consumption and then differentiate (16) totally. We assume that nonforest income is constant and that the subjective rate of time preference equals the rate of discount. The computation in Appendix 1 leads to

$$t_1^* = t_1^*(\overset{\pm}{p}, \overset{+}{\dot{p}}, \overset{\pm}{w}, \overset{+}{m}, \overset{+}{a_0}, \overset{\pm}{\rho}, \overset{-}{T}). \tag{17}$$

Assuming the rotation period does not exceed the period that maximizes the "mean annual increment", it follows that the longer the rotation period, the greater the harvested volume of timber and the greater the average annual timber supply.[4] Thus the comparative statics signs for timber supply are the same as those given in (17).

The effects of timber price and planting costs are ambiguous because of nonconstant price and the presence of both income and substitution effects. However, an upward time trend in price lengthens the rotation period while decreasing price shortens it. Higher nonforest income and higher initial level of nonforest assets imply lower marginal value of consumption

[4] We do not have information on rotation ages in our sample, but the average harvest rate for all observations is 2.81 $m^2$/ha/year, and for harvests above zero 4.21 $m^2$/ha/year. The maximum gross yield of forest stands, i.e., maximum average growth over an infinite time horizon in Southern Finland varies between 5 and 8 $m^3$/ha/year.

and longer rotations. The effect of the rate of interest is ambiguous. Its effect on the marginal value of consumption depends on whether the owner is a borrower or lender. This in turn depends e.g. on whether the stand is cut in the beginning or later in the owners life cycle. In addition, an increase in interest costs has the normal negative effect on the length of the rotation period and finally the *in situ* value of the new stand evaluated at the cutting moment decreases, which lengthens the rotation period.

The rotation period also depends on the owner's age. One reason for this follows from the fact that the length of the period after the stand is cut depends on owner's age. For younger owners, this period has more weight, which decreases the rotation period. However, another effect follows if the subjective time preference $\delta$ differs from the rate of interest $\rho$. If $\delta < \rho$ the marginal value of consumption decreases over time, implying that lengthening the rotation period for enjoying more *in situ* benefits is less costly to older than to younger forest owners. The case $\delta > \rho$ implies the reverse. These effects are demonstrated in Figure 1, which shows a numerical example for the dependence of timber supply on forest owner's age [5].

As we already noted, price fluctuations have the effect that there may be multiple locally optimal moments for harvesting a given stand. Our comparative static results reveal that increasing the price postpones the optimal harvesting moment and *vice versa.* This has the implication that the optimal harvesting moments may concentrate around those years when the price level is highest. This is demonstrated in Figure 2 where the dotted line shows the development of average stumpage price net of logging and transportation costs over our sample period. The solid lines show the present value utility $W_1$, given that stands with initial ages varying between 70 and 130 years are harvested at the given date. Parameter values are the same as in Figure 1. It is optimal to harvest a stand when the solid line obtains its maximum value. As the example demonstrates, the optimal years for harvest concentrate around the years when the price path obtains local maxima.

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[5] It is assumed that $F(x) = rx(1 - x / K), U(c) = (c^{(1-\alpha)} - 1) / (1-\alpha), A(x) = Ax^{\beta}$, and $K = 500$, $r = 0.048, \delta = 0.035, \rho = 0.035, w = 4000, p = 170, A = 0.05, x_0 = 10, \alpha = \beta = 0.5, m = 10^5$, $a_0 = 15000$ and $T = 30$.

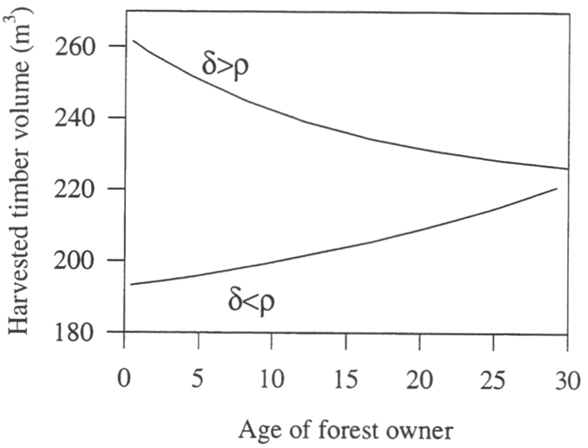


Figure 1. Harvested volume as a function of owner's age

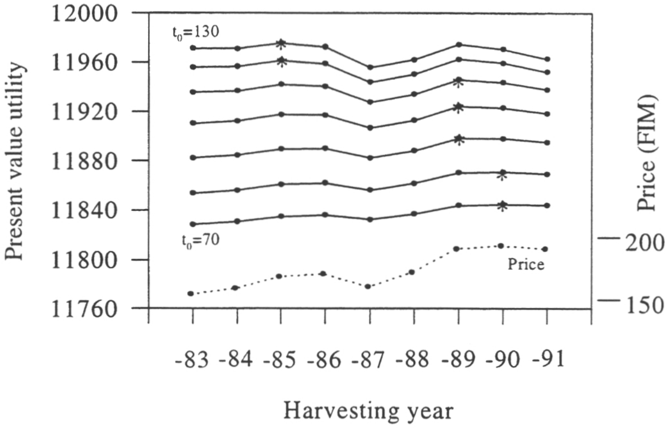


Figure 2. Price fluctuations and optimal timing of harvest

Except for timber price, which in the theoretical model was allowed to have intertemporal variation, the sign predictions of the Faustmann model concern parametric changes in the exogenous variables of the model. Therefore, in order to be able to test them, we must empirically distinguish between possible intertemporal effects and parametric effects of the exogenous variables.

We apply an empirical method developed in labor economics by specifying a timber supply function similar to the " $\lambda$ - constant" life cycle labor supply function. (Heckman and MaCurdy 1981, MaCurdy 1981). The virtue of a $\lambda$ - constant labor supply function is that, given the shadow price of consumption, $\lambda$, the current labor supply depends only on current prices and wages. Thus, the intertemporal substitution effect of wage rate, assuming similar additive preferences accross individuals, is independent of owner-specific variables. The initial value of the shadow price, constant $\lambda$, captures the effects of all other owner-specific information relevant for labor supply decisions. Thus the inclusion of an individual constant term provides a significant reduction in data requirements, while allowing for an unbiased estimate for the intertemporal substitution effect.

The problem of timber supply is more complex than the problem of life cycle labor supply, due to the dynamics of the resource stock. In timber supply function (17), derived using a utility-based rotation model, the parametric effects of prices include both substitution and income effects and depend on the values of all the exogenous variables of the model during the whole planning period. Furthermore, even assuming similar preferences, the intertemporal effect depends also on the shadow price and the timber volume of the stand, in addition to the current timber price. However, this intertemporal effect is unambiguously positive, while the intertemporal effect of current nonforestry income is zero. This provides a basis for testing the model by empirically distinguishing between the intertemporal and parametric effects of the rotation model.

As a first approximation, we specify the intertemporal effects on current harvest rates to be independent on owner-specific factors, given the recent harvesting decisions as indicated by change in timber stock in the previous year.[6] The empirical timber supply function is a sum of two linear functions. The first part, denoted $h_t(.)$, contains the intertemporal effects and explains the annual variation in harvest. The behavioral timber supply equation in (17) is derived under perfect foresight. We assume that the current values of exogenous variables contain all relevant information concerning the intertemporal short-term variation of prices and

[6] Relaxing the assumption of common intertemporal effects would require estimating a random coefficients model with censored dependent variable.

nonforestry income. The second part of the empirical supply function, denoted $\beta_{0i}$, gives the effects of the time invariant observable and unobservable owner characteristics and price and income expectations for the whole planning period. Therefore, the supply of timber of owner i in year t, $h_{ti}$, is written as

$$h_{ti} = h_t(p_{ti}, m_{ti}, \rho_{ti}\, \Delta v_{t-1,i}) + \beta_{0i}\,(p_i^e, m_i^e, v_{0i}, z_i). \qquad (18)$$

Function $h_t$ gives the reaction of harvest to intertemporal changes in prices, $p_{ti}$, incomes, $m_{ti}$, and the interest rate, $\rho_{ti}$. We have assumed perfect capital markets, implying that only the present value of the exogenous income affects the optimal time of harvest. However, it can be expected that the existence of capital market imperfections may cause also the intertemporal changes in nonforest income to affect the harvest rates. Thus, we also include the intertemporal effect of income to test whether the perfect capital market hypothesis is justified. The interest rate should actually be included in the owner-specific part of the function, but because our sample includes only intertemporal variation concerning this variable, it is included explicitly only in function $h_t$. Thus, an important unobserved owner-specific variable affecting the optimal time of harvest is the interest rate.

According to our theoretical model, forest owners harvest stands. In our empirical sample we do not have information on ages of harvested stands. A small holding owner may harvest only a few times or not at all over a reasonable length of a panel data. The more the owner has harvested during the recent past, the more likely he does not presently have economically mature stands, i.e. the age of the remaining stands is under the optimal rotation age. This means that the empirical data should display negative autocorrelation between present and past harvest. However, lagged harvest may be zero. Therefore, lagged difference in the timber stock, $\Delta v_{t-1,i}$, has more information, as it also shows an increase in the physical potential to harvest.

Function $\beta_{0i}$ is an individual constant term that captures the effects of the time-invariant owner characteristics and expectations. In particular, parametric changes in expected timber price, $p_i^e$, nonforestry income, $m_i^e$, initial timber stock, $v_{0i}$, and in other observed and unobserved owner-specific variables, $z_i$, affect the optimal rotation when *in situ* preferences are present. This part of the supply function is assumed to be constant during the study period for an individual forest owner but to vary across owners. Therefore, it represents the effects on harvest rate of unexpected parametric changes in the exogenous variables.

## 2.2 Estimable timber supply function

The theoretical model in section 2.1 led to an implicit behavioral supply relation, which was specified as a sum of two linear functions in equation (18). We further specify the estimable equation for common slope parameters as a linear function of the intertemporal variation of the relevant exogenous variables,

$$h_{ti} = \beta_{0i} + \beta_1 p_{tj} + \beta_2 m_{ti} + \beta_3 \rho_{ti} + \phi \Delta v_{t-1,i} + \varepsilon_{it}, \qquad (19)$$

where, $\Delta v_{t-1,i} = v_{t-1,i} - v_{t-2,i}$, $\phi > 0$, and $\varepsilon_{it}$ is an identically and independently distributed error term. By not including the level of timber stock we avoid the possible strong multicollinearity between it and the time-invariant individual constant term. On average, only 3 to 5 per cent of the stock is harvested annually, and in the empirical data the standard deviation of timber stock across indiviuals is 41.1 while across individuals and time 40.3 (mean 130.9$m^3$/ha).

The not-directly observable individual constant term, $\beta_{0i}$, is specified as a linear function of the owner characteristics, expected values of timber prices, exogenous nonforestry income, and initial timber stock. The equation for $\beta_{0i}$ is written as

$$\beta_{0i} = \theta_0 + \theta_1 v_{0i} + \theta_2 p_{0j} + \theta_3 m_{0i3} + \theta_4 z_{0i} + e_i. \qquad (20)$$

The error term, $e_i$, is normally and independently distributed with zero mean and constant variance. Because timber price and nonforestry income are independent of harvest decisions, their averages can be used to measure expectations without introducing simultaneity in the specification. The observation period average timber stock cannot be used because it is determined simultaneously with the harvest decisions. Therefore, the predetermined initial timber stock is used instead. The age of the forest owner is assumed to be part of the individual effect (z variables in equation (20)). Thus, the age of the owner may represents the taste difference between different generations of forest owners, although theoretically there can also be a life cycle variation in the harvest.

As we do not have all the information on the variables affecting the individual effects, equations (19) and (20) must be estimated in two stages. Initially, equation (19) is estimated in order to obtain the common slope parameters and the individual constant term for each owner. In the second stage, the estimated values for individual effects are regressed against the time-invariant owner characteristics and proxies for expected prices and income. The total effect of the exogenous variables on harvest is the combined effect of these two components.

# 3 Data

We use a sample of 119 farmers from survey data collected by personal interviews in 1985, 1986 and 1991 in Southern Finland. The sampling procedure was two-stage areal cluster sampling, where a farm's probability of entering the sample was proportional to its total land area. In personal interviews the response rate was 94%. Annual micro data on the quantities harvested during the five years preceding the interview were collected, giving annual harvest rates for each owner during the period 1982-1991. For each wood lot, the timber stock in 1982, 1985 or 1986 and 1991 was measured and the stock at the beginning of each year was calculated using the recursion equation, $v_{ti} = (1+g_i)v_{t-1,i} - h_{ti}$. The average percentage growth, $g_i$, was iterated using linear growth and information on the actual stock in 1982, 1985/1986, and 1991 and the actual harvests in each year. In the sample each owner had harvested at least once during the period studied. The annual timber prices are regional prices for pine saw logs in the 47 communities where the sampled wood lots were located. The data was augmented with annual observations on forest owners' income based on tax information. All monetary units are in FIM, deflated by the 1990 wholesale price index.

# 4. Empirical results

## 4.1 Econometric specifications

In this chapter, we first demonstrate the problem with the theoretical interpretation of the results from models using pooled time-series cross-section data, neglecting the unobserved owner characteristics. Then we estimate the interpretable common slope parameters, i.e. the intertemporal effects, of the supply equation (19) using individual dummy variable models. Finally, the cross-section variation providing the parametric effects of the rotion model is studied using the estimated individual effects. The latter results are compared to those that could have been obtained directly using the means of study period harvest and initial (1982) timber stock.

We estimate a model for the expected annual harvest by specifying following empirical equation:

$$h_{ti}{*} = h_{ti} = \beta_i{'}X_{ti} + \varepsilon_{ti} \text{ if the right hand side is positive} \qquad (21)$$

$$h_{ti} = 0 \text{ otherwise,}$$

indicating a censored dependent variable (e.g. Maddala 1983). However, for comparision, we also provide the results of a sales probability model, where the dependent variable is a dummy variable (1=$h_{ti}$> 0, 0). Vector $\beta_i{'} = [\beta_{0i}, \beta]$ includes individual constant terms and the common slope parameters, and $\varepsilon_{ti}$ is the IID error term. Exogenous variables, $X_{ti}$, are as in $h_t(.)$, equation (19).

Ignoring the effects of relevant unobservable variables may create an omitted variables bias in the estimates of the common slopes. The usual way to model unobserved owner characteristics is to use fixed or random effects models. In the present case, a fixed-effects model must be used because the individual effects are known by the forest owner when making harvesting decisions and are correlated with the other exogenous variables of the model.

A consistent estimator for the fixed-effect logit model was proposed by Chamberlain (1980). In the case of a probit or tobit model, no consistent parametric estimator for fixed-effects specification exists (e.g. Hsiao 1989). However, Heckman (1981) demonstrated that the bias in a probit model with a finite sample may not be large, if lagged endogenous variables are not included (cf. Cassel et al. 1996).[7] The dependent variable of the tobit model contains more information than the dependent variable of the probit model. We examined the finite sample bias in the estimates of the common slopes using artificially created data with N=100 and T=10 (Appendix 2). According to the results of 30 simulated samples, the bias in the common slopes of the fixed-effects tobit model is not severe when the distributions of the dependent and independent variables were similar to those in the actual data, even though the error term is not normal. In the base case, which corresponds closest to the actual data, the coefficient of timber price is biased 2 % and the coefficient of income 1 %.[8] The false null, $\beta = 0$, was accepted for prices in 13 and for income in 70 per cent of the samples. However, it was accepted almost exclusively when using pooled data. The bias in the estimates of the parametric effects was more severe, 10 % for the price and 15 % for income, but even this does not give reason to doubt the qualitative implications of the results to be presented.

## 4.2 Intertemporal effects

In order to demonstrate the problem of ignoring the effects of unobserved owner characteristics, we first report the results of a pooled tobit model without allowing for unobserved owner characteristics (table 1, column 2). In this model, neither the effect of timber price nor exogenous income is statistically significant according to the asymptotic standard errors. Therefore, besides the negative effect of the interest rate, the expected harvest seems to be affected only by the variables not included in the traditional optimal rotation model.

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[7] Cassel et al. (1996) estimated a fixed-effects probit model for blue collar worker absenteeism, but in their case T was very large. Thus they could assume the incidental parameter problem to be a small.

[8] The bias increases with the random variation in the individual effects. In the base case the error term of the simulated fixed effect was NID(0,1). The mean of the simulated fixed effects was -3.5 and standard deviation 2.6. The mean of the fixed effects was estimated to be -5.7 and standard deviation 2.8. The corresponding mean and standard deviation estimated from the actual data are -4.6 and 2.9. Using a NID(0,5) error term when simulating fixed effects increased the bias in the intertemporal price effect to 4% and in the intertemporal icome effect to 8%.

In earlier studies using pooled data, price difference or present and lagged prices have been used to estimate intertemporal price effects. The problem with the former is that price differences do not contain any information on the cross-sectional variation. Therefore, the hypotheses derived from the Faustmann model cannot be tested. The problem with the latter is that present and lagged prices move very close to each other, resulting in imprecise estimates, especially when T is small. Aronsson (1990) used price expectation based on opinions of interviewed owners. This type of information is not available in our sample.

Before going to the correct way of estimating the intertemporal effects, we tentatively investigate the source of the effects in the pooled estimation. Table 1, columns 4 and 5 report the results of a tobit model where the annual observations on harvest during 1983-1991 are explained using study period averages of the explanatory variables and initial timber stock measured in year 1982. This eliminates the intertemporal variation from the exogenous variables except for the interest rate, which has only intertemporal variation in our sample. The closer to each other the elasticities of models I and II, the more dominant the effect of cross-section variation in the sample.[9] According to the results, the implied elasticities for 1982 stock, forest owners age and forest land area are close to those obtained using annual observations for exogenous variables. Thus, the marginal effects of the pooled tobit model of these variables are mainly due to cross-sectional variation. Finally, timber price has a negative, not-statistically-significant coefficient in model II, and the absolute value of the not significant negative coefficient of nonforestry income is halved compared to model I.

These experiments suggest that in order to accurately measure the intertemporal effects of prices and income the unobservable owner characteristics must be taken into account. We do this by using individual effects models.[10] The fixed-effects logit model for probability of harvesting is reported in column 2 of table 2. Because Chamberlain's specification does not

[9] Even if data contain both cross-section and over-time variation, they are not necessarily correlated with the corresponding variation in the explained variable.

[10] The individual effects are strongly correlated with initial timber stock and owners age. If these are included on the left hand side of the fixed effects tobit model, correlation between estimated fixed effect and timber stock is over -0.9, with owners age over -0.7. The absolute values of the elasticities became very large (age - 12.8, stock 8.1), indicating a misspecification, as confirmed by the Monte Carlo experiments.

Table 1. Tobit model for annual harvest, 1983-1991, pooled data for 119 NIPF owners. Elasticities are calculated at the means.

| Independent variables, model I / model II | tobit model I, coefficients (standard errors) | elasticities, all observations | tobit model II, coefficients (standard errors) | elasticities, all observations |
|---|---|---|---|---|
| Constant | -2.5615 (2.9215) | | 13.883 (8.8954) | |
| Lagged timber stock/ 82 stock | 0.0146 (0.0047) | 0.392 | 0.0149 (0.0050) | 0.363 |
| Price /average 1983-91 | 0.0173 (0.0128) | 0.830 | -0.0543 (0.0378) | -2.601 |
| Income /average 1983-91 | -0.0043 (0.0029) | -0.097 | -0.002 (0.0032) | -0.035 |
| Interest rate | -0.1954 (0.0617) | -0.278 | -0.1911 (0.0596) | -0.194 |
| Age of the owner/average age | -0.0447 (0.0147) | -0.494 | -0.0462 (0.0165) | -0.467 |
| Has forestry plan | 1.2452 (0.4252) | | 1.0543 (0.4353) | |
| Has bank loans | 0.9229 (0.3985) | | 0.9645 (0.4025) | |
| Forest land/average forest land | 0.0137 (0.0039) | 0.203 | 0.01412 (0.00389) | 0.209 |
| $\sigma$ | 5.3741 (0.1509) | | 5.3867 (0.1511) | |
| N | 1071 | | 1071 | |
| Loglikelihood | -2482.610 | | -2483.192 | |
| $F(\beta'X/\sigma)$ | 0.6075 | | 0.6067 | |

provide estimates for individual constant terms, marginal effects and elasticities cannot be calculated.

According to the Hausman test, the existence of individual effects, even in the model for probability of entering the market, cannot be rejected. The positive effect of price on the probability to harvest is due to intertemporal variation in prices and can be interpreted using figure 2. More forest owners are in the market when prices are high than when they are low. Also, intertemporal variation in nonforestry income affects the probability of harvesting. This is contrary to the perfect capital market hypothesis, which says that only the present value of future income should affects the harvest. The interest rate has a negative

Table 2. Estimation results for fixed-effects logit model for the probability of harvest and for a fixed-effects tobit model for annual per hectare harvest.

| | Fixed effect logit | Fixed-effects tobit model | |
|---|---|---|---|
| Independent variables | Coefficients (standard errors) | Coefficients ($\beta$) (standard errors) | Elasticities for all observations |
| Fixed effect | | -4.4606<br>(2.9227) | |
| $\Delta$ Stock | 0.0121<br>(0.0130) | 0.1192<br>(0.0291) | |
| Price | 0.0074<br>(0.0076) | 0.0371<br>(0.0131) | 1.927 |
| Income | -0.0051<br>(0.0025) | -0.0125<br>(0.0054) | -0.306 |
| Interest rate | -0.0427<br>(0.0406) | -0.2209<br>(0.0586) | -0.339 |
| Dummy 1885 | -0.6422<br>(0.2969) | | |
| $\sigma$ | | 4.9372<br>(0.1383) | |
| Exp. hrv. $m^3/y$ | | 2.74 (0<4.49) | |
| N | 1071 | 1071 | |
| Log likelihood | -361.65 | -2392.196 | |
| Autocorrelation | | -0.11 | |
| $\Phi$ ($\beta$'X/$\sigma$) | | 0.612 | |
| Hausman statistics | 25.68(4) | 40.3(4) | |

coefficient but is not statistically significant. In the two-period modelling tradition, also this has been interpreted to indicate an imperfect capital market with effective credit rationing (cf. Koskela 1989, Kuuluvainen and Salo 1991).

A dummy variable was included in the fixed-effects logit model for sales probability to take into account a structural change. In the mid-1980s, the average size of positive sales increased, while the number of owners annually in the market decreased. In 1983-1985 the average sale above zero was 3.76 $m^3$/ha/year, in 1986-1991 the respective figure was 4.71, while the average sale for all observations was only 0.2 $m^3$/ha/year larger during 1986-1991 than 1983-1985. The reason may be in the development of harvesting technology and increased share of stumpage sales due to a decreased markup for delivered wood. However, also the capital market liberalization completed in 1986 may be one reason for the development. Under imperfect capital markets, forest owners have an incentive to smooth harvest income over time

(Tahvonen, 1997). Similar incentive does not exist in a perfect capital market, as the harvest income can be allocated to optimal consumption over time through borrowing and lending.

In the fixed-effects tobit model, the signs of the coefficients are the same as in the model of probability of harvesting, with relatively small standard errors. The Hausman test of the fixed-effects specification against the random-effects tobit model (not reported) favors the fixed-effects model as expected.[11] Also, the absolute value of the loglikelihood of the fixed-effects model is clearly smaller than that of the tobit model using pooled data. The coefficient of the lagged difference in timber stock is positive and significant, suporting the hypothesis of negative autocorrelation in harvest.[12] The elasticity of expected harvest with respect to stumpage price is 1.9 for all observations. Also the absolute value of the elasticity of income is larger than in the pooled estimation. These results and our Monte Carlo experiments suggest, that in the present case pooled estimation underestimates the intertemporal effects.

To conclude, according to the fixed-effects tobit model, timber prices are the key determinant of intertemporal variation in timber supply. Aslo, intertemporal variation in nonforestry income and the interest rate affect annual variation in the quantity harvested. The question that remains open is whether prices and income are connected to variation in harvest between forest owners.

### 4.3 Cross section effects

In empirical studies the effects of parametric changes are normally interpreted to represent the long run-effects. They can in principle be measured either by using cross section data from a particular year or averages of the variables over some longer period of time. However, one may not be able to extract the cross section variation by regressing e.g. average harvest during

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[11] In the random-effects model, the effect of price was negative, the effect of income positive, but neither was significant. Timber stock had a negative coeffficient and the age of the owner a positive coefficient. The latter are hard to interpret and may be due to the correlation between the individual effects, timber stock and age.

[12] The absolute size of the coefficient does not have a clear interpretion and may be biased (Hsiao, 1989) as also indicated by the Monte Carlo experiments. However, when the corresponding variable was excluded from the model in Monte Carlo simulations, the price effect became underestimated by 16% and the income effect by 13 %.

the study period on means of exogenous variables. This is due to the fact that a forest owner with a low income may be supplying on average the same amount of timber as a forest owner with high nonforestry income. This is because the location of the supply curve may be different because of unknown owner-specific factors. By using the estimated fixed-effects, more information can be obtained. Even if two forest owners with different average income may have the same average timber supply during the period studied, the individual constant terms representing the unknown individual owner characteristics can differ. If the expected income level or some other relevant variable affects the harvest, this relationship can be extracted by regressing the individual constant terms on the average income along with other measurable determinants of timber supply.

To study the cross section variation, the estimated individual effects from the tobit model are regressed against the time-invariant owner characteristics and the means of the study period stumpage prices, nonforestry income and the start-of-1982 timber stock. The parameter estimates of this regression are consistent and asymptotically normal assuming that the number of observations per person over time is large (Heckman and MaCurdy 1980; see also appendix 2). We also compare the two-stage estimation results to the information contained in the timber stock and average harvest. Therefore, table 3 reports OLS models for average harvest and 1982 timber stock with the same explanatory variables, along with the linear regression model for estimated fixed-effects.

The fixed-effects model's standard errors indicate that all the exogenous variables have a statistically significant effect on the estimated owner-specific constant term of the tobit model (the dummy variable for bank loans only at the 10 per cent level).[13] According to the $R^2$ of this model, one-third of the variation in individual effects can be explained by observable owner characteristics and proxies for expectations. In the model explaining the 1982 timber stock, nonforestry income, whether the owner had aquired bank loans in 1986-91, owners age,

[13] In addition to the variables reported in table 3, we experimented with a number of other owner characteristics (agricultural land area, owner's place of residence, principal use of the farm, number of children, bequest motive), but the t-values remained under 0.5 and the F-test suggested rejection of the variables from the regression.

Table 3. Regression models for estimated fixed-effects, 1982 timber stock and average harvest for 1983-1991 (standard errors are in parentheses).

| Independent variables | Linear Regression model for fixed-effects (θ) | Linear Regression model for timber stock in 1982 | Linear regression model for average harvest |
|---|---|---|---|
| Standing stock 1982[1)] | 0.01564 | | 0.0129 |
| | (0.0064) | | (0.0039) |
| Timber price (47 observ.) | -0.0962 | 0.7986 | -0.02706 |
| | (0.0480) | (0.8531) | (0.0353) |
| Nonforestry income | 0.0093 | 0.1446 | -0.0007 |
| | (0.0041) | (0.0605) | (0.0025) |
| Forest land area | 0.0157 | -0.1427 | 0.0050 |
| | (0.0049) | (0.0740) | (0.0031) |
| Age of owner | -0.0519 | 0.5257 | -0.0222 |
| | (0.0188) | (0.2657) | (0.0112) |
| Has forestry plan[2)] | 1.0753 | 0.1241 | 0.3534 |
| | (0.5397) | (7.9004) | (0.3254) |
| Has bank loans[2)] | 0.8747 | -17.962 | 0.7446 |
| | (0.5117) | (7.3914) | (0.3125) |
| Forestry Board district 1 | | 9.8089 | -0.1967 |
| | | (10.179) | (0.4210) |
| Forestry Board district 2 | | -12.253 | -0.2681 |
| | | (8.986) | (0.3739) |
| Constant | 15.439 | -92.767 | 8.0866 |
| | (11.439) | (202.91) | (8.3653) |
| N | 119 | 119 | 119 |
| $R^2$ | 0.35 | 0.20 | 0.21 |
| F-test | (7,114) 8.53 | (8,113) 3.52 | (9,112) 3.23 |
| Mean fixed effect /stock/harvest | -4.643 | 119.054 | 2.805 |

1) Standing timber stock at start of 1982, fixed effect and averages of the other dependent variables are calculated using observations from 1983-1991.
2) Has forestry plan and Has bank loans are dummy variables.

and forest land are statistically significant.[14] In the equation for average harvest, 1982 timber stock, owners age, and whether the owner had aquired bank loans in 1986-91 had statistically significant coefficients at the 5 per cent level.

[14] Using the study period initial stock as the dependent variable means assuming that forest owners base their harvesting decsions on expected long-run values of the relevant exogenous variables. The assumption is justified by the data. Regressing end-of-study-period timber stock on the means of the exogenous variables produces the same number of statistically significant coeffcients with slightly higher t-values and implied absolute values for the elasticities. We also note that in Finland NIPF owners have relatively good information on their forest. In the sample, 75 % of owners had a 10-year forest management plan. Even those owners who

The negative effect of the stumpage price on individual effects implies that supply tends to be lower for forest owners perceiving high average prices than for forest owners facing low average prices, indicating a backward bending long-run supply curve. The negative sign is predicted by the basic Faustman model. The sign is ambiguous when nontimber benefis are included, but the substitution effect seems to dominate with realistic parameterizations. The positive effect of the average income indicates that wealthier forest owners have a tendency to supply more than less wealthy forest owners. This is consistent also with the statistically significant positive coefficient of nonforestry income in the model for 1982 timber stock. Wealthy forest owners can afford more nontimber benefits than less wealthy forest owners. The Negative coefficient for age means that older generations of NIPF owners harvest less cubic meters per hectare per year than younger ones.[15]

Monte Carlo experiments in appendix 2 indicate that we can draw tentative conclusions concerning the total effects of prices and income on timber supply. In a linear model, $\hat{\beta}_b = \hat{\theta} + \hat{\beta}_{cv}$ (Mundlak 1978, Hsiao 1988, Laisney et al. 1992), where $\hat{\beta}_b$ is the parameter for the between-group variation, $\hat{\theta}$ is the estimated parameter from the model explaining fixed-effects using the means of the exogenous variables of the fixed-effects model, and $\hat{\beta}_{cv}$ is the estimated parameter of the fixed-effects model. Subsequently, the total parametric effects at the means can be calculated as a sum of the common slopes and the parameter estimates from the linear regression model for fixed effects.[16] The marginal effects of the tobit model take into account the probability to harvest, and the total effect can be calculated using $\hat{\beta}_b = \Phi(\hat{\beta}'X/\hat{\sigma})(\hat{\theta}+\hat{\beta}_{cv})$. Thus we treat the parameters of the linear regression model explaining fixed effects as if they had been estimated simultaneously with the common slopes.

do not have a written plan normally use a professional extention forester's help in planning harvests and silvicultural investments.

[15] The slightly larger elasticity of harvest with respect to age of owner in the pooled tobit model than in the tobit model using averages of the exogenous variables in table 1 indicates that there may be variation in harvest also over an individual forest owner's lifetime.

[16] MaCurdy (1981) calculated the effects of permanent and temporary changes in wage rate the labor supply in a similar manner. In his model the elasticity of supply with respect to wage rate is a pure substitution effect, while the parameters of the model explaining individual effects give the wealth effects.

The elasticity of harvest with respect to a permanent price difference between forest owners is -3.1. Following MaCurdy (1981) this can be interpreted as an elasticity of timber supply with respect to a permanent parametric change in price. The corresponding elasticity calculated direcly from the model explaining the average harvest (column 4 in table 3) is -2.3, but the coefficient is not significant. The tobit model with annual observations for harvest but with study period averages of the explanatory variables gives an elasticity of -2.6 (t-value of the coefficient is 1.43). Even though these are not statistically significant, the values of the elasticities lead to similar qualitative conclusions.

An example of a permanent unexpected price change affecting supply might be a change in the sales tax rate. Another relevant elasticity from the empirical point of view is the elasticity due to a temporary parametric change. Following MaCurdy (1981) we can compute this by using an average ownership time for a woodlot, which in Finland is 30 years. The elasticity of timber supply with respect to a temporary parametric change can be calculated by dividing the effect of a permanent change by the average time for adjusting the harvest and adding this to the intertemporal effect, i.e. $\hat{\beta}_b = \Phi(\hat{\beta}'X/\hat{\sigma})(\hat{\theta}/15+\hat{\beta}_{cv})$, which gives the elasticity of 1.6. Thus, the intertemporal effect dominates the effect of a temporary change in the stumpage price. The elasticity of a permanent change in exognous income is negative but small in absolute terms, -0.08. The effect of a temporary change (- 0.3) is clearly dominated by the negative elasticity of intertemporal change.

Although the absolute values of the elasticities strictly describe only the present data set, they suggest the usefulness of panel data in testing theoretical hypotheses concerning timber supply. In particular, cross section data cannot be used to model the short-term dynamics, which are important e.g., when NIPF owner behavior over business cycles is of interest. On the other hand, using time series data, parametric changes that are important in testing the theoretical hypothesis of the Faustman rotation model may be difficult to capture with aggregate time series data unless long observation period allowing for the methods of cointergration analysis are available.

# 5 Conclusions

This paper demonstrates a method for investigating intertemporal and parametric (i.e. long-run) effects of exogenous variables on timber supply, using cross-section time series micro data. The effects of the (expected) intertemporal changes in the exogenous variables are important when forest owner behavior, e.g. over a business cycle is of interest. However, when the implications of policy changes are considered, the effects of permanent or temporary parametric (unexpected) shifts in variables affecting the rotation age should be known. The absolute values of the elasticities of the present study cannot be generalized, but we hope that the results would motivate empirical research and collection of panel information which has a long tradition in other fields of economics such as consumer behavior and labor economics.

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## Appendix 1

We cannot rule out multiple local maximum, but by the second order necessary conditions it must hold that $\partial\Gamma / \partial t_1 < 0$ in any given interior maximum,

$$\text{where } \Gamma = p(t_1-)\lambda(t_1)\{F[x(t_1-)] - x(t_1-)[\rho - \dot{p}(t_1-)/p(t_1-)] + \rho w / p(t_1-)\}$$
$$+A[x(t_1-)] - A(x_{t_1}) - \varphi(t_1)F(x_{t_1}) = 0.$$

The assumption $\delta = \rho$ implies by (7) that consumption is constant. Assuming also that nonforest income is constant, yields by (2) and (3):

$$a(t) = [a_0 + (m-c)/\rho]e^{\rho t} + (c-m)/\rho, \text{ for } t \in [0, t_1) \text{ and} \tag{A1}$$

$$a(t) = \left[(m-c)e^{-\delta T}/\rho\right]e^{\rho t} + (c-m)/\rho, \text{ for } t \in [t_1, T]. \tag{A2}$$

The fact that $a(t_1) = a(t_{1^-}) + px(t_{1^-}) - w$ yields, using (A1) and (A2), that

$$c = \rho\left[px(t_1-) - w + a_0 e^{\rho t_1}\right] / \left[e^{\rho t_1} - e^{\rho(t_1 - T)}\right] + m. \tag{A3}$$

We thus obtain:

$$\partial c / \partial p = \rho x(t_1-) / \mu > 0, \tag{A4}$$

$$\partial c / \partial w = \rho x(t_1-) < 0, \tag{A5}$$

$$\partial c / \partial m = 1 > 0, \tag{A6}$$

$$\partial c / \partial a_0 = \rho e^{\rho t_1} / \mu > 0, \tag{A7}$$

$$\partial c / \partial \rho = \left\{\mu - \rho\left[t_1 e^{\rho t_1} - (t_1 - T)e^{\rho(t_1 - T)}\right]\right\}\left[px(t_1-) - w + a_0 e^{\rho t_1}\right] / \mu^2 + \rho^2 a_0 e^{\rho t_1} \overset{>}{\underset{<}{=}} 0, \tag{A8}$$

$$\partial c / \partial T = -\rho^2 e^{\rho(T - t_1)}\left[a_0 e^{\rho t_1} + px(t_1-) - w\right] / \left(e^{\rho T} - 1\right) < 0, \tag{A9}$$

where $\mu = e^{\rho t_1} - e^{\rho(t_1 - T)} > 0$. Because $\partial\lambda / \partial c < 0$ and $\partial\Gamma / \partial t_1 < 0$, we obtain

$$\partial\Gamma / \partial p = \partial\lambda / \partial p Y + \lambda\left\{F\left[x(t_1-)\right] - \rho x(t_1-)\right\} \overset{>}{\underset{<}{=}} 0 \quad \Rightarrow \frac{dt_1}{dp} \overset{>}{\underset{<}{=}} 0, \tag{A10}$$

$$\partial\Gamma / \partial \dot{p} = \lambda x(t_1-) / p(t_1-) > 0 \quad \Rightarrow \frac{dt_1}{d\dot{p}} > 0, \tag{A11}$$

$$\partial\Gamma / \partial w = \partial\lambda / \partial w Y + \lambda\rho \overset{>}{\underset{<}{=}} 0 \quad \Rightarrow \frac{dt_1}{dw} \overset{>}{\underset{<}{=}} 0, \tag{A12}$$

$$\partial\Gamma / \partial m = \partial\lambda / \partial m Y > 0 \quad \Rightarrow \frac{dt_1}{dm} > 0, \tag{A13}$$

$$\partial\Gamma / \partial a_0 = \partial\lambda / \partial a_0 Y > 0 \quad \Rightarrow \frac{dt_1}{da_0} > 0, \tag{A14}$$

$$\partial\Gamma / \partial \rho = \partial\lambda / \partial\rho Y - \lambda\left[px(t_1-) - w\right] - \partial\varphi(t_1) / \partial\rho \overset{>}{\underset{<}{=}} 0 \quad \Rightarrow \frac{dt_1}{d\rho} \overset{>}{\underset{<}{=}} 0, \tag{A15}$$

$$\partial\Gamma / \partial T = \partial L / \partial T Y - \partial\varphi(t_1) / \partial T < 0 \quad \Rightarrow \frac{dt_1}{dT} < 0, \tag{A16}$$

where $Y = p(t_1-)\left\{F\left[x(t_1-)\right] - x(t_1-)\left[\rho - \dot{p}(t_1-) / p(t_1-)\right] + \rho w / p(t_1-)\right\} < 0.$

## Appendix 2

In this appendix we study the possible bias in the estimates of the common slope parameters of the dummy variables tobit model for harvest, using simulated data. Also the information in the estimated individual constant terms, i.e. incidental parameters, is briefly discussed. Models with fixed effects produce biased MLE estimates in finite samples (cf. Hsiao 1989). When the estimators of common slopes are not independent of the incidental parameters, the bias in the incidental parameters tends to be carried over to the common slopes (cf. Heckman 1981, Hsiao 1989). Therefore, we compare the bias caused by fixed effects in a nonlinear MLE of the tobit model and the omitted variables bias when a pooled sample is used.

The data used in the Monte Carlo experiments was created with Limdep 7.0 (Green 1995). The end-year per hectare timber stock, $V_{t,i}$, is calculated using the recursion formula, where the end-of-year stock is the beginning-of-year stock plus the growth generated by a logistic growth function and minus harvest during the year, i.e.

$$V_{t,i} = V_{t-1,i} + g(V_{t-1,i}) - H_{t,i}, \qquad \text{(B1)}$$

where g is $0.06(V_{t-1,i} - (V_{t-1,i})^2/400)$ and $H_{t,i}$ harvest during the year. The linear index for harvest, HI, is calculated as follows,

$$HI_{ti} = \beta_{0i} + \beta_1 P_{tj} + \beta_2 M_{ti} + \beta_3 HI_{t-1,i} + \varepsilon_{ti}, \qquad \text{(B2)}$$

where the $\beta_k$, k=1,2,3, are common slope coefficients, $\beta_{0i}$ is the individual constant term, the Ps represent regional prices (j=1,2,3,..34) and Ms exogenous income (i=1,..100), t is time (t=0,.. 12) and the random term, $\varepsilon_{ti}$, is independently and indentically distributed. The observed harvest is

$$H_{ti} = HI_{ti} \text{ if } HI_{ti} > 0 \text{ and zero otherwise.} \qquad \text{(B2')}$$

Thus there is no direct effect from level of the annual beginning-of-period standing timber stock on harvest, but the lagged index creates a negative autocorrelation in the dependent variable. A positive harvest during the previous year reduces the probability of harvest and the quantity, if the index is positive. A negative value for the index in the previous period increases the probability of harvest and also the quantity.

Individual effects were assumed to depend on age, $age_{1i}$, initial standing stock, $V_{0i}$, the initial district price, $P_{1j}$, and forest owner's initial exogenous income, $M_{1i}$, and a normally distributed random term, $\varepsilon_{0i}$ ~NID(0, $\sigma$), as follows:

$$\beta_{0i} = \theta_1 (V_{0i}) + \theta_2 P_{0j} + \theta_3 M_{0j} + \theta_4 age_{0i} + \varepsilon_{0i}, \qquad \text{(B3)}$$

where $\theta_k$, k=2,3,4, are constant coefficients and $\theta_1$ is a logistic function. The individuals do not correct the harvesting "plan" during the simulation according to the actual stock. However, the estimation results from the actual sample presented in table 1, indicate that adding the annual variation in timber stock does not change the elasticity of harvest with respect to the stock, and the distribution of actual harvest can be replicated by using this specification.

Average timber price was modelled as a trend stationary AR-1 process, $P_t = 1.0*t + 0.98*P_{t-1} + NID(0,10)$. Regional prices are calculated using $P_{tj} = d_j*P_t + NID(0,1)$, where $d_j$ is a constant between 0.95 and 1.05. This roughly corresponds to the behavior of actual timber prices (due to the nationwide recommended price agreement system during the period). The exogenous income is also an AR-1 process, $M_{ti}=1.25*t+0.95*M_{t-1,i}+\varepsilon_{ti}$ , where $\varepsilon_{ti}$ is a uniform random variable with mean zero and range –25 to 25 (cf. Heckman 1981, p. 189). The parameters for the price and income equations were initially estimated from the actual data. The initial values for price, income and age were generated using normally distributed (pseudo) random numbers.

The initial timber stock is a function of initial values of timber price, forest owner's nonforestry income and age, $V_{0i} = 20 + 0.5*P_{0i} + 0.1*M_{0i} + 0.4*age_{0i} + NID(0,40)$, which

Table A1. Descriptive statistics for stock and harvest in the actual data (N=1071) and for the respective variables and individual effects in artificial data (N=1000).

| | Mean | Std. Dev | Skew. | Kurt. | Minimum | Maximum | N |
|---|---|---|---|---|---|---|---|
| P | 161.5 | 14.3 | 0.6 | 2.4 | 136.7 | 204.2 | 1000 |
| Price | 166.5 | 12.4 | 0.3 | 1.7 | 146.3 | 187.0 | 1071 |
| M[1)] | 146.4 | 53.1 | 0.1 | 2.4 | 20,4 | 294.9 | 1000 |
| Income | 110.0 | 67.7 | 1.3 | 5.1 | 3.8 | 448.8 | 1071 |
| AGE | 40.6 | 9.5 | 0.4 | 2.7 | 22 | 70.3 | 1000 |
| Age | 54.7 | 13.9 | 0.1 | 2.0 | 22 | 88.0 | 1071 |
| V | 137.4 | 40.4 | 0.1 | 2.7 | 47.3 | 253.4 | 1000 |
| Stock | 130.9 | 41.1 | 0.6 | 3.3 | 49.2 | 294.4 | 1071 |
| H | 2.5 | 3.8 | 2.4 | 10.6 | 0 | 28.3 | 1000 |
| Harvest | 2.8 | 4.0 | 2.5 | 10.3 | 0 | 25.7 | 1071 |
| $\beta_0$ | -3.5 | 2.6 | -0.5 | 3.0 | -10.8 | 1.7 | 1000 |
| $\beta_0$(est)[2)] | -5.7 | 2.7 | | | -11.2 | -0.04 | 1000 |
| $\beta_0$(est)[3)] | -4.6 | 2.9 | | | -13.7 | 0.04 | 1071 |

1) The difference is due to 10 wealthy owners in the actual data.
2) Estimated using simulated data. Average of 30 simulations
3) Estimated using actual data

again roughly corresponds to the parameter estimates from the actual data.[17] However, lower bounds for timber stock and age were introduced in order to guarantee that the owners did not begin planning under age zero and/or with negative timber stock.

The data was generated (for t=0,..12) and the models estimated (for t=3,..12) 30 times (except for the price, for which the same series were used in repeated simulations, due to the space limitation of Limdep 7.0) using a nonnormal error term in harvesting index.[18] Table A1 gives the average descriptive statistics for the simulated data (15 first samples used in the estimations) and the descriptive statistics for the actual data. To create a dependent variable with the same distribution as the harvest in the actual data, the error term in the linear index is generated by normal distribution NID(0,2.5) with positive values raised to the power 1.6. Figure A1 plots the histograms of actual and simulated harvest. The distribution of the actual harvest cannot be reproduced using the normal error term even if negative autocorrelation is introduced. The share of nonzero observations in the simulated data is 57 %, in the actual data

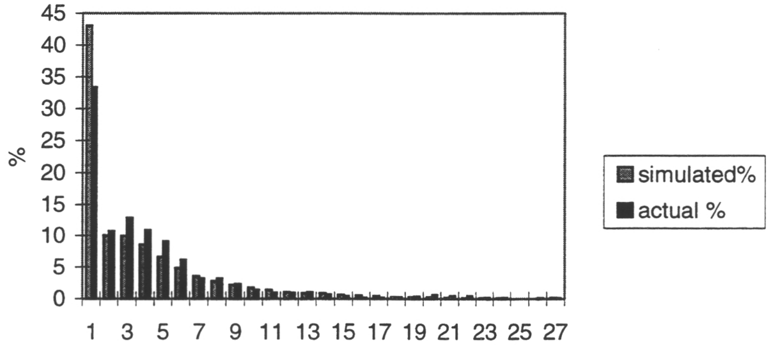


Figure A1. Distribution of actual harvest and sampling distribution of means of 15 samples of simulated harvest in table A1: h=0, 0<h≤1, 1<h≤2 etc., where h is harvest in cubic meters per hectare per year.

[17] This relation is needed to make the data imitate the actual data, because we simulate the data only 3 periods before the 10-period sample for estimations. Thus the dependence of initial stock on prices and owner's income and age makes the data behave as if the process had been going on for some time.

[18] Seed for the random number generator was 4444 4444.

66 %. However, the results are not sensitive to censoring, at least if it is not much over 50 per cent.

Table A2 presents the results of the Monte Carlo experiments. Pooled and fixed-effects tobit models in columns 2-5 are estimated using the data generating process of Table A1 and two different model specifications. In the first case, I, the pooled and fixed-effects models are intentionally misspecified. The pooled tobit model is estimated without age as an exogenous variable (column 2). Compared to true intertemporal effects of the DGP, the estimated parameters are biased. The t-values (not reported) accept the false null hypothesis, $\beta = 0$, for these variables in almost all cases. Only the effect of lagged timber stock captures the cross-sectional variation.

Ignoring the source of variation may lead to difficulties also in a fixed-effects model. In the second case, where a fixed-effects tobit model is used (I, column 3), the age of the owner is assumed to have an intertemporal effect on harvest and also the lagged timber stock is included

Table A2. Results of Monte Carlo estimations using simulated data.[1)] The figures in parentheses are standard errors of the coefficients in 30 simulations.

| | I, NID(0,1) | | II, NID(0,1) | | III, NID(0,5) | | IV, NID(0,10) | |
|---|---|---|---|---|---|---|---|---|
| Parameters | tobit model | fixed-effects tobit model | tobit model | fixed-effects tobit model | tobit model | fixed-effects tobit model | tobit model | fixed-effects tobit model |
| P | 0.0004 | 0.0367 | 0.0061 | 0.0377 | 0.0152 | 0.0353 | 0.0202 | 0.0257 |
| | (0.0111) | (0.0155) | (0.0111) | (0.0156) | (0.0158) | (0.0164) | (0.0245) | (0.0151) |
| M | –0.0024 | -0.0113 | –0.0022 | -0.0119 | 0.002 | -0.0113 | 0.0077 | -0.0108 |
| | (0.0038) | (0.0084) | (0.0036) | (0.0091) | (0.0078) | (0.0916) | (0.0123) | (0.0092) |
| $V_{(t-1)}$ | 0.0450 | 0.4377 | 0.04671 | | 0.0105 | | -0.0525 | |
| | (0.0068) | (0.0496) | (0.0074) | | (0.0094) | | (0.0144) | |
| Age | | -1.1170 | -0.0553 | | -0.0407 | | -0.0019 | |
| | | (0.1423) | (0.0193) | | (0.0441) | | (0.0797) | |
| $\Delta V_{(t-1)}$ | | | | 0.6106 | | 0.5287 | | 0.3922 |
| | | | | (0.0702) | | (0.0534) | | (0.0433) |
| const | -5.3197 | -15.915 | –4.2680 | -5.7416 | -1.8350 | -5.4298 | 3.9943 | -4.0401 |
| | (2.147) | (3.637) | (2.2382) | (2.7673) | (3.1494) | (2.8244) | (4.4829) | (3.3895) |

1) Number of simulations is 30. N=100 and T=10. The linear index is $HI_{ti} = \beta_{0i} + 0.037{*}P_{tj} - 0.012{*}M_{ti} - 0.35{*}HI_{t-1.i} + IID(0.9,4.2)$. and the equation for fixed effect is $\beta_{0i} = 0.14{*}(V_{0i} - V_{0i}^2/400) - 0.1{*}P_{0j} + 0.01{*}M_{0j} - 0.05{*}age_{0i} + NID(0,\sigma)$.

in the model, even if it does not appear in the linear index for harvest. The intertemporal price and income effects are only slightly biased. However, the elasticities implied by the coefficients for timber stock and age cannot be interpreted in the light of the data generating process. The implied elasticity of harvest with respect to initial stock is over 13 while the implied elasticity with respect to age is about -10. When a similar specification was used with the actual data, the elastcity of harvest with respect to beginning-of-period stock was 8.1, with respect to age – 12.8. The elasticity of harvest with respect to price also became unrealistically large.

Next, all the variables that are in the data generating process are included in the pooled tobit estimation for harvest (II, column 4). This does not have a marked effect on the coefficients for price and exogenous income. The coefficient estimated for age is 10% higher than its coefficient in the DGP. The coefficient for the timber stock is somewhat higher than in the previous pooled estimation representing the effect of the initial stock on harvest through the fixed effect. Thus, even when all variables of the data generating process are included, the interpretation of a pooled estimation requires caution.

Finally, the fixed-effects tobit model with the same specification as is used with the actual data for intertemporal variation in annual harvest is estimated (II, column 5). The upward bias in the coefficient for timber price is 2 % while the coefficient for income is biased downward by 1 %. Thus, the fixed-effects tobit model can produce relatively unbiased estimates if the model is correctly specified and censoring is not very strong. However, the random variation in the actual individual effects is not known. Increasing the standard deviation of the random term of the fixed effects in DGP from 1 to 5, increases the bias in the estimated intertemporal coefficient for timber price to 4 % and in the coefficient for income to 8 %. If the standard error of the random term is further increased to 10, the bias in the price coefficient becomes 30 % and in the income coefficient 10 %. However, in this case the pooled tobit model produces a wrong sign prediction for the effect of income and a negative coefficient for the timber stock. These results contradict the results obtained from the pooled estimation using the actual data, and this DGP is obviously not a good approximation of the actual data.

If the process is allowed to continue, e.g. for ten years before sampling without updating the fixed effects, the bias in the common slopes increases. In this case, the fixed-effects tobit model (specification II, in table A2) underestimates the price coefficient by 12 % and overestimates the income coefficient by 8%. However, it is probably not unrealistic to assume that forest owners update their information affecting individual differences in harvest rate every 5 to 10 years. Therefore, the conclusion is that even if there may be random variation in individual effects, the fixed-effects model is more likely to lead to correct qualitative conclusions than the pooled estimation.

With the present DGP, the random effects tobit model cannot be used for Monte Carlo simulations, because it is very difficult to get convergence, even if the individual effects are made purely random. Covergence was also a problem with the actual data. As there is theoretical justification and empirical evidence for the correlation between the exogenous variables and the individual effects, only the results of the fixed-effects tobit model are reported (see footnote 5 of the main body of the text). In addition, in the light of our theoretical model, the interpretation of the random effects model is just as difficult as the interpretation of the pooled tobit model.

Heckman (1981) pointed out that a problem with the fixed-effects probit estimator was the fact that the false null, $\beta = 0$, was accepted more often than for the random effects probit model. In Heckman's (1981) data, the individual effects were not dependent on the exogenous variables of the model, and therefore the fixed-effects model was actually not supported. In the present case, standard t-statistics based on the asymptotic covariance matrix of the fixed-effects tobit model (II, table A2) accepted the false null for prices in about 13 % and for income in 70% of the samples of the simulated data. Using pooled data, the false null was accepted exclusevly for prices and in 27 samples out of 30 for income. Poor efficiency, rather than bias, is the problem of the fixed-effects tobit model in the present case. Especially, the standard errors of the coefficient of income are large in table A2. This suggests caution when interpreting the absolute values of the estimates from the actual data.

Finally, as we make use of the estimated fixed effects in the main text, it is of interest to see how well the variation in actual individual constant terms is captured by the estimated fixed effects. Table A3 reports results from this experiment. Model I (column two) represents the best unbiased estimates of the relation between true DGP fixed effects and the estimation period averages of the exogenous variables. Model II (column three) gives the corresponding estimates as measured by the estimated fixed effects. The average $R^2$ for this model, (II), in five first samples is 0.26. This compares to 0.35 for the corresponding model estimated using actual

Table A3. Results of Monte Carlo experiments for simulated fixed effects[1)] in DGP and corresponding fixed effects estimated using simulated data and the tobit model (II) in table A2. The figures in parentheses are standard errors of the coefficients in 30 simulations.

| Parameters | I[2)] | II[3)] | III[4)] | IV[5)] |
|---|---|---|---|---|
| P | -0.0877 | -0.0965 | -0.0568 | -0.0444 |
| | (0.0247) | (0.0543) | (0.172) | (0.2008) |
| M | 0.0103 | 0.0118 | 0.0246 | 0.0268 |
| | (0.0025) | (0.0111) | (0.0189) | (0.0234) |
| V | 0.0500 | 0.04358 | -0.0723 | -0.0914 |
| | (0.0043) | (0.0094) | (0.0163) | (0.0210) |
| Age | -0.0600 | -0.0651 | -0.0177 | -0.0031 |
| | (0.0114) | (0.0299) | (0.1055) | (1.2309) |
| Constant | 4.7436 | 4.7152 | 11.989 | 11.1123 |
| | (4.0573) | (8.9838) | (28.417) | (31.616) |

1) Number of simulations is 30, N=100, the linear index is $HI_{ti}$ as above, and the equation for fixed effect is $\beta_{0i} = 0.14^* (V_{0i} - V_{0i}^2/400) - 0.1^*P_{0j} + 0.01^*M_{0j} - 0.05^*age_{0i} + NID(0,\sigma)$.
2) Actual simulated fixed effects against means, $\sigma$=1.
3) Fitted fixed effects against means, $\sigma$=1
4) Actual simulated fixed effects against means, $\sigma$=10
5) Fitted fixed effects against means, $\sigma$=10

data. In model II the bias compared to model I where the true simulated fixed effect is used is 10 % for price and 15% for income. In Model III the dependent variable is the true simulated fixed effect of model IV in table A2. In model IV the dependent variable is the estimated fixed effect from model IV in table A2. The two latter simulations (III, IV) demonstrate, that if fixed effects are burried under to large random variations (or specification is not justified), the estimates from the second stage are unreliable, as indicated by the large standard errors of the coefficients. The t-values (not reported) of model IV are close to zero as are the $R^2$s. According to table A2, fixed-effects specification for estimating common slopes is still a better choice than pooled estimation. To conclude, incidental variables bias the estimates of cross-sectional variation more than they bias the estimates of the common slopes. However, in spite of the poor efficiency, the qualitative conclusions based on the estimated parameters of the two step estimation process are correct in the simulated data.

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