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Intergenerational income immobility in Finland: contrasting roles for parental earnings and family income

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Abstract An intergenerational model is developed, nesting heritable earning abilities and credit constraints limiting human capital investments in children. Estimates on a large, Finnish data panel indicate very low transmission from parental earnings, suggesting that the parameter of inherited earning ability is tiny. Family income, particularly during the phase of educating children, is shown to be much more important in shaping children's lifetime earnings. This influence of parental incomes on children's earnings rises as the children age because the returns to education rise. Despite Finland's well-developed welfare state, persistence in economic status across generations is much higher than previously thought.

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Omena ei putoa kauas puusta. (The apple doesn't fall far from the tree).
Syntynyt kultalusikka suussa. (Born with a golden spoon in his mouth).

Traditional Finnish adages.

Three distinct explanations have been postulated for the commonly observed, positive correlation between incomes of children and those of their parents. One hypothesis rests upon the intergenerational transmission, either genetically or through the home environment, of unobserved earning abilities. This variant suggests an auto-regressive process in earnings across generations. A second approach emphasizes budget constraints, limiting parents' investments in human capital of their children. A third alternative models parental preferences with respect to their children as endogenous to their own parents' incomes, resulting in a three-generation auto-regressive process.[1]

Parental resources may potentially shape economic success of children via a number of channels, in addition to the duration of education, such as through the quality of education, extracurricular opportunities, neighborhood influences, and postschool training (Restuccia and Urrutia 2004). Yet, in comparison to the numerous analyses of schooling achievement outcomes,[2] attempts to discern the relative contributions of auto-regression in earning ability versus family incomes in the process of lifetime earnings' formation remain few. Mulligan (1997, 1999) offers an important exception, estimating little difference in the intergenerational regression toward the mean in wages between US adult children who have (or expect to) inherit more than $25,000 relative to those who do not anticipate such bequests, supporting Mulligan's conclusion, "...that the observed intergenerational dynamics of measures of economic status are not the result of borrowing constraints" Mulligan (1999, p. S215). However, Gaviria (2002) finds Mulligan's result to be sensitive to the specific inheritance break-point imposed and reaches the opposite conclusion. Corak and Heisz (1999) suggest that a slower regression toward mean earnings among the Canadian sons of middle-income families, than among their wealthier counterparts, may reflect liquidity constraints on middle-class parents. Grawe (2004) disagrees and applies a quantile regression approach to the same data and finds no evidence to support the importance of liquidity constraints. Björklund et al. (2005) examine an extraordinary Swedish data set, which permits comparisons in earnings of several sibling types, indicating a dominant role for a genetic component rather than an environmental effect.

[1]Becker and Tomes (1979, 1986), Behrman et al. (1995). Easterlin et al. (1980) expressed the third alternative in terms of fertility outcomes, though it may be applied equally to quality rather than just quantity of children. Excellent surveys are provided in Behrman (1997), Solon (1999), Björklund and Jäntti (2009), Björklund and Salvanes (2010), and Black and Devereux (2010).

[2]Contributions include Cameron and Heckman (1998, 2001), Behrman and Rosenzweig (2002), Carneiro and Heckman (2002), Plug and Vijverberg (2003, 2005), Black et al. (2005), Belley and Lochner (2007), Lochner and Monge-Naranjo (2011), and Pronzato (2012).

Yet, as Björklund et al. emphasize, identifying the effects of nature versus nurture from birth-type comparisons requires some strong assumptions.[3]

In Section 1, a model nesting both intergenerational auto-regression in earning ability and a family budget constraint is outlined. The nested model is estimated, drawing upon a very remarkable, and largely unexplored, data panel encompassing the entire Finnish population from 1970 to 1999. The transition to an empirical, regression model, related prior work and some general issues in estimation are considered in Section 2. The data, and prior work on intergenerational transmission in Finland, are described in Section 3. Initial estimates for both sons and daughters, as well as some sensitivity tests, follow in Section 4. Both inherited ability and the influence of family budget constraints may, in principle, shape earnings of the next generation, and this is recognized in the nested model. However, all of our estimates point to small and fairly insignificant values for the Becker–Tomes parameter of inherited earning ability, no matter whether the father, mother, or both are considered as sources. Yet the results suggest throughout that family incomes are an important factor shaping intergenerational transmission in Finland. Section 4 also includes some evidence, from a second data set, on the influence of both parents' and grandparents' incomes on earnings in the third generation. Following Behrman and Taubman (1989), these estimates permit a test of the endogenous preference hypothesis of Easterlin et al. (1980).

Previous evidence on Finland, and on the Nordic countries more generally, suggests significantly greater intergenerational mobility than in the USA or UK.[4] Cross-country comparisons of estimates are fraught with differences in specification and data (Solon 2002; Corak 2006; Bratsberg et al. 2007; Jäntti et al. 2007). Nonetheless, the estimates of intergenerational transmission elasticities presented here approximate some of the existing estimates for the USA and UK. This suggests that, although Finland clearly has a more equal distribution of income than the USA, mean intergenerational mobility across the spectrum in Finland may be quite limited.

In the US context, sons' earnings appear to become more like those of their fathers as the son ages (Reville RT, Intertemporal and life cycle variation in measured intergenerational earnings mobility, unpublished). Section 5 confirms a significant, parallel tendency among both sons' and daughters' earnings in the Finnish context. Haider and Solon (2006) note that this property may derive from age dependence in the association between observed and lifetime earnings. In Section 5, a specific form of this association is investigated, stemming from rising returns to education with age. Supporting results on the age dependence of returns to education in Finland, including cross-sibling comparisons, are also presented in Section 5. After accounting for the interaction between age and the returns to education, no further significant

[3] See also Duncan et al. (2005), Sacerdote (2007), Liu and Zeng (2009), and Behrman et al. (2011).

[4] See, for example, Gustaffson (1994), Björklund and Jäntti (1997), Österbacka (2001), Björklund et al. (2002), Jäntti et al. (2007).

rise in the intergenerational transmission elasticity is observed, as Finnish sons and daughters age.

The seminal contribution of Becker and Tomes (1986, p. 10) incorporates, "...the effect of imperfect access to debt contracted for children." There remains the possibility of credit constraints on parental borrowing against their own future income, and hence the effect of timing of income receipts upon intergenerational transmission. In Section 6, it is demonstrated that parental income when the child is of school age, even after instrumenting for these income measures, has a larger impact on the child's subsequent earnings than does family income thereafter.[5]

Section 7 closes by attempting to put these fresh results in perspective. Whether lessons may be derived directly for other countries from these results on Finland remains to be seen; similar nested models appear not to have been estimated elsewhere. Nonetheless, it is compelling to note the importance of family incomes, even in this Finnish context where the welfare state has attained such importance.

1 Models of intergenerational transmission: with and without credit constraints

The Becker and Tomes (1979, 1986) model of intergenerational mobility is in the class of consensus, parental preference models, in which children play no role in decision making (Behrman 1997). Mulligan (1999) lends specificity to this framework by assuming a single child family in which the utility (U_{p}) of the parent takes a CES form:

$$U_{\mathrm{p}} = \left[C_{\mathrm{p}}^{\eta} + \alpha\ Y_{\mathrm{c}}^{\eta}\right]^{1/\eta} \tag{1}$$

where C_{p} represents consumption of the parent, Y_{c} is permanent income of the child, α is the weight placed on altruism, and $(1-\eta)^{-1}$ is the elasticity of substitution. The parent chooses an amount (H_{c}) to invest in the child's human capital, bequests and inter vivos transfers (B_{c}) to be passed to the child, and the parent's own consumption level, so as to maximize Eq. 1 subject to two constraints. The first is a budget constraint:

$$C_{\mathrm{p}} + H_{\mathrm{c}} + B_{\mathrm{c}} \leq Y_{\mathrm{p}} \tag{2}$$

where Y_{p} represents wealth of the parent. In the second constraint, each child's wealth is generated from the child's human capital and transfers such that

$$Y_{\mathrm{c}} = G_{\mathrm{c}} H_{\mathrm{c}}^{\rho} + B_{\mathrm{c}}\,(1+\iota) \tag{3}$$

[5]In a related paper, Oreopoulos et al. (2008) show that earnings are lower among a group of Canadians, whose fathers lost their jobs in plant closings, than among a measurably comparable group whose fathers were not retrenched. See also Eide and Showalter (2000).

where G_c denotes inherent capacities of the child to generate earnings, ρ is the return on human capital (with $0 < \rho < 1$), and ι is the rate of interest obtained on transfers. In the event that an interior solution exists for both H_c and B_c, the first-order conditions may be solved to provide

$$H_c = \left[\rho G_c (1 + \iota)^{-1}\right]^{1/(1-\rho)} . \tag{4}$$

Provided that ι is independent of the amount of bequest, the budget constraint is linear and the parent's choice of H_c is independent of the elasticity of substitution; hence, η does not appear in Eq. 4. The parent simply equates the returns on investment in the child's human capital with that on bequests (see, however, Jürges 2000).

On the other hand, negative transfers (borrowing against the child's wealth at the same rate ι) may well be infeasible (Becker and Tomes 1986). If credit constraints restrict transfers to be non-negative, then the optimal choice of the parent is given from the first-order condition on maximizing Eq. 1 with respect to H_c, subject to Eqs. 2 and 3 and the additional constraint that $B_c \geq 0$. When this last constraint is binding, the solution has the implicit form:

$$Y_p = H_c + \left[\rho\alpha\ G_c^{\eta}\ H_c^{\rho\eta-1}\right]^{1/(\eta-1)} \tag{5}$$

in which $\partial H_c / \partial Y_p > 0$, given G_c. In the instance of a Cobb–Douglas utility function, Eq. 5 reduces to the more tractable, explicit form:

$$H_c = Y_p\ \rho\alpha\,(1 + \rho\alpha)^{-1} . \tag{6}$$

For convenience, the case represented by Eq. 4 is referred to as the wealth model in the remainder of this paper, while Eq. 5 is termed the credit-constrained case.[6] A critical distinction may be drawn between these two cases. In the wealth model, investments in the child's human capital are affected only by the child's inherent earning capacity G_c and, in particular, parental wealth does not appear in Eq. 4. In contrast, in the credit-constrained case, parental wealth indeed affects the child's human capital directly in both Eqs. 5 and 6. Thus, in the credit-constrained case, parental wealth already impacts that of the child through the budget constraint, but in the wealth model, another factor must be introduced to generate an auto-regressive process in earnings across generations. In Becker and Tomes (1979, 1986), this additional factor is a presumption of auto-regression in inherent earnings capacity, akin to the seminal ideas of Galton (1869), embodying an unknown mix of genetic and cultural transmissions, which may be written here as

$$G_c = G_p^{\psi}\gamma_c \tag{7}$$

where G_p is the inherent earning capacity of the parent, γ_c is a stochastic component in this inheritance, and $0 < \psi < 1$.

[6]These correspond to the high-resource and low-resource family cases in Behrman et al. (1995).

Let E_c represent that component of the child's wealth which we may call lifetime earnings, such that

$$E_c = G_c\ H_c^{\rho}. \tag{8}$$

In the case of the wealth model, one can then substitute in Eq. 8 for H_c from Eq. 4 to provide

$$E_c = \left[(1+\iota)\,\rho^{-1}\right]^{\rho/(\rho-1)} G_c^{1/(1-\rho)}. \tag{9}$$

If investments in the parent's human capital were also dictated by the wealth model, then a relationship equivalent to Eq. 9 would hold for the parent's earnings too and may be inverted as

$$G_p = E_p^{1-\rho^*}\left[\left(1+\iota^*\right)^{-1}\rho^*\right]^{-\rho^*} \tag{10}$$

where G_p, E_p, ι*, and ρ* correspond in the parent's generation to G_c, E_c, ι, and ρ in the case of the child. From Eqs. 10, 7, and 4, investment in the child's human capital in the wealth model is then given by

$$H_c = \left\{E_p^{\psi(1-\rho^*)}\left[\rho\gamma_c\,(1+\iota)^{-1}\right]\left[\left(1+\iota^*\right)^{-1}\rho^*\right]^{-\psi\rho^*}\right\}^{1/(1-\rho)}. \tag{11}$$

Moreover, combining Eqs. 7, 9, and 10 yields the intergenerational, auto-regressive earnings structure in this wealth model

$$E_c = \kappa_0\ E_p^{\psi(1-\rho^*)/(1-\rho)}\gamma_c^{1/(1-\rho)} \tag{12}$$

in which κ_0 depends only upon the parameters ι, ρ, ι*, ρ*, and ψ. Note the implication that, if the returns on human capital remain approximately constant across generations (that is $\rho \approx \rho$*), then the elasticity of the child's lifetime earnings with respect to earnings of the parent simply represents the elasticity of transmission of inherent earnings ability, ψ, in this model.

In the case of the wealth model, a pure income effect does not change investments in the child's human capital, so Y_p does not affect the child's earnings in Eq. 12. The child's wealth (Y_c) increases with Y_p, even in this case, but this is entirely through bequests rather than enhanced earnings. On the other hand, more able parents with higher earnings invest more in the child's human capital, as in Eq. 11. Positive auto-regression in intergenerational earnings derives from the direct effect of inherited ability upon the child's earnings and from the additional human capital of the more able child.

For any credit-constrained family with homothetic preferences, investment in human capital expands with wealth, even when earning ability is not inherited ($\psi = 0$). Whether human capital investments are amplified if $\psi > 0$ depends upon the shape of preferences for these credit-constrained families.[7] In the case of a Cobb–Douglas utility function, human capital investments are

[7] In Eq. 5, the partial derivative of H_c with respect to G_c has the same sign as η, thus being negative (positive) for elasticities of substitution <1 (or >1).

independent of G_c and hence independent of inherent ability transmission, as may be seen from Eq. 6. Nonetheless, the child's earnings remain enhanced directly by inherent abilities in Eq. 8, so combining Eq. 8 with Eqs. 6, 7, and 10 provides

$$E_c = \kappa_1 \, Y_p^{\rho} \, E_p^{\psi(1-\rho^*)} \gamma_c \tag{13}$$

where κ_1 is a function of the parameters α, ρ, ι^*, ρ^*, and ψ. In Eq. 13, parental wealth enhances the child's earnings through a budget constraint effect. Moreover the parent's earnings are correlated with parental ability and hence the child's ability and earnings, provided $\psi > 0$.

2 Empirical counterparts

Permanent incomes are unobservable and data on entire lifetime earnings are nowhere available to date. Estimation of equations such as Eq. 12 or 13 must rely, instead, upon observed incomes or earnings of individuals and of their parents, tracked over some limited duration within the life cycle. For the moment, a simple, quite traditional approach is adopted, to be extended in Section 5. Let the natural logarithm of child c's earnings observed at time t (e_{ct}) be written

$$e_{ct} = \lambda_c + \mathbf{\Lambda} \, \mathbf{A}_{ct} + \xi_{ct} \tag{14}$$

where λ_c is a fixed effect for person c, $\mathbf{A}_{ct}$ is a vector of polynomial terms in c's age at time t (a_{ct}), $\mathbf{\Lambda}$ is a vector of coefficients, and ξ_{ct} is a stochastic disturbance term. Similarly, let the logarithms of earnings and income for parent p (of child c), observed at time τ, be given respectively by

$$e_{p\tau} = \lambda_p + \mathbf{\Lambda}^* \, \mathbf{A}_{p\tau} + \xi_{p\tau} \tag{15i}$$

$$y_{p\tau} = \pi_p + \mathbf{\Pi}^* \, \mathbf{A}_{p\tau} + \upsilon_{p\tau} \tag{15ii}$$

where $\mathbf{A}_{p\tau}$ is a vector of polynomial terms in age of parent p at time τ, λ_p and π_p denote personal fixed effects, $\mathbf{\Lambda}^*$ and $\mathbf{\Pi}^*$ are vectors of parameters, and $\xi_{p\tau}$ and $\upsilon_{p\tau}$ are stochastic disturbance terms.

Since (for now) age profiles are assumed common across individuals in Eqs. 14 and 15, the logarithms of the lifetime earnings and permanent income terms, E_c, E_p, and Y_p, may be defined to equal the individual fixed effects λ_c, λ_p, and π_p plus terms common across observations. Substituting in Eqs. 12 and 13, respectively, in the wealth model and credit-constrained cases leaves

$$e_{ct} = \omega_0 + \omega_1 \, e_{p\tau} + \Omega \, \mathbf{A}_{ct} + \Omega^* \mathbf{A}_{p\tau} + \mu_{ct} \tag{16}$$

$$e_{ct} = \chi_0 + \chi_1 \, e_{p\tau} + \chi_2 \, y_{p\tau} + X \, \mathbf{A}_{ct} + X^* \, \mathbf{A}_{p\tau} + \nu_{ct}. \tag{17}$$

Both models are nested within

$$e_{ct} = \beta_0 + \beta_1 \, e_{p\tau} + \beta_2 \, y_{p\tau} + B \, \mathbf{A}_{ct} + B^* \mathbf{A}_{p\tau} + \varepsilon_{ct} \tag{18}$$

with $\beta_2 = 0$ if the wealth model (Eq. 16) holds. In addition, since $\omega_1 = \psi(1 - \rho^*)(1-\rho)^{-1}$ and $\chi_1 = \psi(1-\rho^*)$, provided $(1-\rho^*) > 0$, the inheritance of earning ability (ψ) approaches zero as β_1 goes to zero. In other words, this offers a key to exploring the role of inherited earning ability as modeled by Becker and Tomes (1979, 1986). Presuming the rate of return to parent's education (ρ*) < 1, if the estimate of β_1 approaches zero, this would indicate an inherited earning ability parameter, ψ in Eq. 7, that also approaches zero.

Issues in estimation In instances where multiple time-period observations on the earnings of the child are available, a number of options have been explored to estimate intergenerational transmission elasticities. In his initial approach, Zimmerman (1992) treats each year of data on sons' earnings in cross section and then considers the mean of these independent estimates. Abul Naga (2002) notes that a more efficient estimator of this mean across time periods is obtained from a between-individual estimator, adopting the mean earnings of the child as dependent variable, provided that the transitory components of the child's and parent's earnings are stationary. Such between-individual estimators are, perhaps, the most common choice (Behrman and Taubman 1990; Mulligan 1999). A third alternative is to take fuller advantage of the panel features of the data.[8] The results presented in Section 4 are derived from between-individual, time-average estimators. However, in the extensions considered in Section 5, panel estimates are also presented.

The disturbance term in Eq. 18 is composed of

$$\varepsilon_{ct} = \rho_0\ \ell n\gamma_c - \beta_1\ \xi_{pt} - \beta_2\ \upsilon_{pt} + \xi_{ct} \tag{19}$$

where $\rho_0 = (1-\rho)^{-1}$ in the wealth model and $\rho_0 = 1$ in the credit-constrained case. In general, ε_{ct} thus includes time-dependent errors in measurement from representing the parent's lifetime earnings or permanent income by a short panel of observations (ξ_{pt} or υ_{pt}). The well-known consequence is that ordinary least squares (OLS), applied to equations such as Eq. 18, generates inconsistent estimates. Two main approaches have been adopted to address this inconsistency.

The first though less common approach to the errors in measurement in Eq. 18 has been to instrument a single year of data on a parent's earnings.[9] The second approach has been called the method of averages, in which parental earnings or incomes are measured by the mean of time-series observations on parents (early applications include Behrman and Taubman (1990), Altonji and Dunn (1991), and Solon (1992)). Abul Naga (2002) considers some properties

[8] For example, Altonji and Dunn (1991) and Zimmerman (1992) both apply GMM estimators to their panel data, while Lillard and Kilburn (1997) maximize a joint likelihood function derived from an ARMA structure in transitory earnings.

[9] Exploratory results in the present context suggest that the use of instrumental variables increases the estimated transmission from family income but not from parental earnings, though identifying appropriate instruments is, as always, a difficult task and this route is not followed in the present paper.

of such a method of averages estimator in a wealth model, in which $\beta_2 = 0$ and the age terms ($\mathbf{A}_{ct}$ and $\mathbf{A}_{p\tau}$) are suppressed for simplicity. Assuming that the three remaining stochastic components, $\ell n\gamma_c$, ξ_{pt} and ξ_{ct}, are stationary, homoskedastic, and mutually uncorrelated, Abul Naga notes that the probability limit for the OLS estimator of β_1 in Eq. 18 is then

$$\text{plim}\ \widehat{\beta_1} = \beta_1 - \frac{\beta_1 \sigma_{\xi p}^2 \Big/ T_p}{(1-\rho^*)^2\, \sigma_{G_p}{}^2 + \sigma_{\xi p}^2 \Big/ T_p} \tag{20}$$

where $\sigma_{G_p}^2$ is the variance of $\ell n G_p$, $\sigma_{\xi p}^2$ is the variance of ξ_{pt}, and T_p represents the number of periods over which the parent's earnings are averaged. The estimator in Eq. 20 is biased toward zero, but the extent of attenuation bias diminishes with T_p. In his survey, Solon (1999) stresses the importance of larger values of T_p and some testing along these lines is reported in Section 4.

3 The Finnish context: data and prior work on Finland

In the mid-1960s, personal identity codes were introduced in Finland. These identity codes enable Statistics Finland to access information on individuals across administrative registers, such as the Central Population Register and Tax Register. Since 1970, Statistics Finland has compiled a population census every 5 years and by 1990 the census was entirely register-based. By matching the unique personal identifiers across the censuses, Statistics Finland has constructed a Longitudinal Census Data File with panel data on the entire population of Finland at 5-year intervals from 1970 to 1995. In addition, since 1987, Statistics Finland has maintained the Longitudinal Employment Statistics file which is updated annually. Since the same personal identifier is adopted in both the Census and the Longitudinal Employment Statistics, the two data sets can be merged, providing panel data on each resident of Finland for 1970, 1975, 1980, 1985, and then annually from 1987 through 1999. Throughout the entire database, cohabiting families are assigned a common family identification number.[10] Thus, it is possible to identify the parent(s) living with a child in 1970 then to trace the child and parent(s) through to 1999.[11]

[10]"A family consists of a married or cohabiting couple and their children living together; or a parent and his or her children living together; or a married or cohabiting couple without children. Persons living in the household-dwelling unit who are not members of the nuclear family are not included in the family population, even if they are related" (Statistics Finland 1995, p. 16).

[11]Whether the adults with whom the child was living in a 1970 nuclear family are the biological parents is not known. Children recorded in the family unit comprise biological and adopted children of either spouse, though foster children and children in the care of the family are not classified as part of the family.

Table 1 Descriptive statistics

Sons and daughters	Sons		Daughters	
Age range in 1970	0–16		0–16	
Number of individuals	32,714		31,504	
Percent with single parent in 1970	7.5		7.7	
Number age 10+ in 1970	14,964		14,543	
Number age 10+ with positive earnings observations	14,394		13,891	
Sons and daughters age 10+ with positive earnings	**Mean**	**SD**	**Mean**	**SD**
Number of positive earnings observations per individual	13.54	3.34	13.26	3.45
Mean age during positive earnings	33.95	2.78	34.03	2.86
Time mean of log earnings	11.33	0.80	10.95	0.80
Time mean ratio of wages to earnings	0.90	0.23	0.93	0.19
Number of family income observations per individual	15.84	3.22	15.94	3.05
Time mean log family income	11.46	0.72	11.47	0.72
Parents	**Fathers**		**Mothers**	
Number of individuals	26,727		28,749	
Number with positive earnings obs.	26,266		26,377	
Percent died by 1999	31.7		12.5	
Percent emigrated by 1999	3.0		3.0	
	Mean	**SD**	**Mean**	**SD**
Age in 1970	39.43	9.94	36.75	9.24
Mean age during positive earnings	52.12	8.03	51.08	7.37
Number of positive earnings observations per individual	9.44	5.60	9.94	5.33
Time mean of log earnings	10.78	1.22	10.39	1.29
Time mean ratio of wages to earnings	0.69	0.39	0.79	0.35

To provide a family-based sample from these data, Statistics Finland drew observations in two stages: the first stage is a 1 %, random sample of individuals from the 1970 Population Census and the second stage comprises all family members cohabiting, at the time of the 1970 census, with the first-stage individuals. For present purposes, observations are drawn from this sample on all children aged 0–16 and living in a family in 1970.[12] This provides panel data on 32,714 sons, 31,504 daughters, and their parents.

Descriptive statistics on this sample are provided in Table 1. Note that nearly 8 % of the sons and daughters are in single-parent families. In addition, almost a third of the initial fathers had died by 1999, though only some 12.5 % of mothers died. The other source of attrition from these register-based, panel data is by emigration, which claimed about 3 % of parents during this interval.

[12] Among the children ages 0–16 in our database, less than three quarters of 1 % are not living with their family and hence omitted from the family-based sample. This very low fraction is in accord with a report by The Population Research Institute (1995), which shows that in the mid-1980s Finnish children continued to live with their parents even beyond age 16 more frequently than did children in the other Nordic countries: 95 % of boys ages 16–19 were still at home in Finland as were 88 % of girls.

Measures of earnings and incomes Estimation of Eq. 18 calls for measures to represent earnings of parents and of their offspring, and to represent family income. Our data include three measures of annual income for each individual: (1) wages and salaries; (2) entrepreneurial (self-employment) income from agriculture, business, and partnerships; and (3) total income subject to state taxation, which includes most social security benefits.[13] The data on wages and salaries for each individual are reported directly to the Government of Finland by employers. The data on self-employment and other taxable income are compiled from the Tax Register. In Finland, taxable income is recorded separately for each person; joint-filing of income taxes does not exist. Throughout, all income measures are expressed in 2000 Finnish Marks, using the annual, national cost-of-living index to deflate.

Two representations of the earnings of sons and of daughters (e_{ct}) are initially adopted as alternative, dependent variables: annual wages (including salaries) and total annual earnings (the sum of wages and self-employment earnings).[14] The same measures on wages and earnings are also available for parents. Self-employment earnings are a more important income source to the parents than to the younger generation: almost a third of the average father's earnings, and about 20 % of the average mother's earnings, are derived from self-employment in years in which they earn (see Table 1). The standard theoretical framework, as outlined in Section 1, refers to the earnings of one parent ($e_{\mathrm{p}\tau}$).[15] Most of the empirical literature assumes this

[13] "Income subject to state taxation does not include scholarships and grants received from the public corporations for study or research, earned income from abroad if the person has worked abroad for at least six months, part of the social security benefits received from the public sector and tax-exempt interest income" (Statistics Finland 1995, p. 18). Since 1985 taxable income in Finland does, however, include child, maternity, and unemployment benefits.

[14] Most estimates of the intergenerational correlation in incomes are based on sample survey data, though a few studies have also extracted income measures from tax registers (Corak and Heisz 1999, on Canada; Österberg 2000, on Sweden; and Mazumder 2005, who uses the US Social Security Administration's Summary Earnings Records). Register-based data, whether derived from employer or tax records, have some advantages over survey responses, particularly where the survey respondent is not the person employed. Yet register-based data also have some limitations. The Finnish data lack information on hours worked, for example, and hence are inadequate to examine hourly rates of pay and preclude computation of full income. The information on self-employment earnings also has some drawbacks. The usual caveats apply with respect to the role of capital income components within self-employment earnings. Moreover, given progressive taxes on each individual, an incentive exists to spread family self-employment income across family members where possible. Nonetheless, self-employment is not a major source of earnings for young people. For only 5.2 % of sons and 5 % of daughters are more than 90 % of their earnings, in an average year, derived from self-employment. For the remaining sons and daughters, self-employment earnings represent only 7 and 4.4 % of their earnings, respectively.

[15] This literature also presumes a single child. An extension to multiple children is to be the subject of a separate paper, though preliminary findings indicate that the results in the present paper prove robust to including controls for family size.

refers to the father.[16] To the extent that such parental earnings are meant to embody the transmission of ability traits from parent to child, through genetic or environmental effects as in Eq. 7, the assumption that fathers are the solitary source may be questioned. Moreover, an exclusive focus on earnings of fathers censors the children of single mothers from the sample (and vice versa). The labor force participation rate of women is high in Finland. For example, in 1980, 78 % of mothers in our sample received earnings and 91 % generated earnings at some stage in the panel. Overall, the time average earnings of mothers represent 42 % of the parental total earnings in two-parent families. In Section 4, a number of options are therefore considered, adopting the earnings of fathers only, of mothers only, including earnings of both mother and father separately, and earnings of whichever parent has the higher mean log earnings.

To represent family income, the taxable incomes accruing to both parents are summed. To continue, such aggregation, even after instances in which parents separate, may not be entirely appropriate. On the other hand, in Finland, payments of alimony and child support are very common following separation, so the continued pooling of both parents' incomes remains eminently reasonable. Although the impacts of parental separation on child outcomes are beyond the scope of the present paper, some results on the sensitivity of intergenerational transmission estimates to alternative definitions of family income, adjusted for separation, are noted in the following section.

In the estimation of the generic form (Eq. 18), the distinction between family income and parent's earnings is critical. In Finland, most transfers from the state are included in the definition of taxable income. Under the Finnish welfare state system, these transfers are a significant component of income for a wide range of families. Our data do not permit a breakdown of unearned income by source. However, separate data, reported by KELA (the social insurance institution of Finland), show transfers, on average across all households in Finland as a percentage of disposable income, rising from about 22 % in 1980 to over 43 % by 1995. Largely as a result of these contributions, there is a substantial variation in the contribution of parental earnings at all levels of family income in our data. For instance, within each percentile of families, ranked by time-averaged family incomes during periods in which the major earner is employed, the interquartile range in the contribution of the father to family income never falls below 29 % of the median contribution, and for mothers, the comparable measure is 37 %. The result is that the mean coefficient of variation in fathers' earnings is nearly 0.5 within each percentile of family income. Moreover, there is a strong negative correlation between combined earnings of the parents and receipt of other income at each level of

[16]Certainly, the majority of empirical studies, both of the USA and elsewhere, seek to relate sons' earnings to earnings of their fathers. A much smaller number of studies correlate the earnings of daughters and of fathers, and a handful of studies have considered the mothers' earnings. See Solon (1999), Tables 3, 4, 5 and 6 for a summary of estimates for both sons and daughters. Subsequent estimates of intergenerational income transmission for daughters include Österbacka (2001) and Chadwick and Solon (2002).

family income. This unearned income falls fairly steadily as a fraction of family income, from about the 4th to the 98th percentile of family income, suggesting the importance of the welfare state system in generating these incomes (for greater detail, see Electronic supplementary material—Appendix A).[17]

Prior work on the Finnish data In a series of papers, Markus Jäntti and Eva Österbacka adopt the quinquennial census portion of the Finnish data to estimate the intergenerational correlation in earnings. Estimates are provided for both sons and daughters, related to either mothers' or fathers' earnings separately. For example, Österbacka (2001) applies ordinary least squares to equations relating the mean annual earnings of sons and of daughters in 1985, 1990, and 1995 to the mean annual earnings of fathers or mothers in 1970 and 1975, controlling for age and age squared of the child and parent. These specifications provide the following estimates of intergenerational earnings transmission elasticities:[18]

	Sons	Daughters
Father's earnings	0.129 (25.8)	0.100 (16.7)
Mother's earnings	0.037 (9.3)	0.023 (4.6)

In comparison to somewhat similar estimates for other countries, Österbacka (2001, p. 480) concludes "that intergenerational mobility is relatively high in Finland." Jäntti et al. (2007) estimate higher transmission elasticities, at least for sons:[19]

	Sons	Daughters
Father's earnings	0.173 (8.9)	0.080 (4.1)
Family income	0.220 (11.1)	0.112 (11.4)

In these latter estimates, earnings observed in 1993 and 2000 are truncated at a lower bound. These measures are then deployed to project what earnings would be at age 40, for sons and daughters born between 1958 and 1960. Fathers' earnings are measured in 1975 when their mean age was 47, again projecting their earnings at age 40 for the right-hand variable. Family income is measured by the sum of both parents' earnings in 1975, at which stage the difference between earnings and total taxable income remained quite small.

[17] The appendices to this paper are available as Boston University Institute for Economic Development Discussion Paper No. 221 at http://www.bu.edu/econ/files/2011/09/Lucas_Kerr.pdf.

[18] Source: Österbacka (2001, Table 3). *t* statistics for a zero null hypothesis are shown in parentheses, calculated from the reported standard errors. In addition, Österbacka explores estimates within quintiles of parents' earnings and the correlations between siblings' earnings, while also undertaking a decomposition of the intergenerational correlations. See also Jäntti and Österbacka (How much of the variance in income can be attributed to family background? Empirical Evidence from Finland, unpublished).

[19] See Tables 2 and 9 in Jäntti et al. (2007). *t* statistics for a zero null hypothesis are shown in parentheses, calculated from the reported confidence intervals.

In large part, these projections and restrictions on the sample are imposed to render results as comparable as possible to those generated by Jäntti et al. for the UK, USA, and other Nordic countries. By these measures, Finland again displays greater intergenerational mobility than the UK and far greater mobility than the USA, though exhibiting a mixed ranking among the Nordic states.

In their study of the impact of the comprehensive school reform enacted in Finland between 1972 and 1977, Pekkarinen et al. (2009) estimate much higher transmission elasticities from fathers' to sons' earnings, on the order of 0.30 prior to the reform and 0.24 after treatment. Here, sons' earnings are measured in 2000, for cohorts born between 1960 and 1966, of which the reform impacted portions of those born from 1961 to 1965 and all born thereafter. Fathers' earnings are represented by the average log of earnings from 1970 through 1990. It is important to note, however, that in this study, the measure of "earnings" for both sons and fathers actually refers to our taxable income measure, rather than wages or wages plus self-employment earnings. Note also that in the following Section 4 the focus is on a set of sons and daughters all of whom were too old to have been affected by this comprehensive school reform.

All of our samples of sons and daughters do, however, fall in an age range that potentially benefitted from the substantial expansion in government guaranteed loans for tertiary-level students, offering subsidized interest rates, which occurred in 1969. Yet, despite this expansion, surveys among students at the time indicate that these government loans and grants represented only about 50 % of subsistence costs (Blomster 2000).

4 Initial estimates

The results in this section are from between-individual regression models that adopt a time-series average of log earnings or wages of sons and daughters as the dependent variable. Initially, only years of positive earnings (wages) are included when computing the mean log earnings (wages) of both child and parent. As a further restriction in this section, only sons and daughters who are at least age 25 in 1985 are included.[20]

Table 2 first reports estimates, obtained by OLS, of transmission from parental earnings (wages) to those of sons and daughters. As mentioned in Section 3, separate estimates are reported with several alternative measures

[20] The age range varies considerably over which earnings of children are included in prior studies. Solon (1992) imposes a lower bound at age 25 in his study of sons' earnings, while the lower bound in Zimmerman (1992) is 29, and Dearden et al. (1997) look at UK sons and daughters when they are 33. Although only children who are at least 25 in 1985 are included in our initial analysis in Section 4, any earlier positive earnings (or wages) of these children are included in computing mean earnings, provided the son or daughter was at least 20 at the time of observation in the 1975 or 1980 census. To the role of age in these estimates, Section 5 returns.

Table 2 Transmission from parental earnings and wages (between-individual estimates on positive earnings)

	Sons					Daughters				
Mean log earnings										
Father's mean log earnings	0.058		0.048			0.044		0.035		
	(10.13)		(7.82)			(7.10)		(5.23)		
Mother's mean log earnings		0.051	0.043				0.055	0.044		
		(9.20)	(7.39)				(9.53)	(7.23)		
Mean log combined earnings				0.062					0.059	
				(11.00)					(9.74)	
Major earner's mean log earnings					0.054					0.055
					(9.96)					(9.55)
No. of observations	12,957	12,528	11,346	14,139	14,139	12,500	12,138	10,993	13,645	13,645
Mean log wages										
Father's mean log wages	0.124		0.096			0.042		0.027		
	(19.61)		(12.72)			(7.71)		(4.07)		
Mother's mean log wages		0.102	0.067				0.051	0.038		
		(14.00)	(8.27)				(8.36)	(5.42)		
Mean log combined wages				0.141					0.055	
				(22.23)					(10.25)	
Major earner's mean log wages					0.144					0.061
					(20.49)					(10.11)
No. of observations	11,691	11,024	9,338	13,377	13,377	11,340	10,732	9,118	12,954	12,954

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix

of parental earnings. The only other explanatory variables included in these specifications are the age of the child and of the parent(s), each expressed as a separate polynomial to the fourth power. More particularly, the ages of both the child and parent are measured as the mean age during which positive earnings are reported.[21] The correspondence of a mother's observed earnings to her lifetime earnings may well differ, at any given age, from that of a father. Consequently, when looking at the major earner's earnings, a dummy for their gender is also included, both separately and interacted with their polynomial age profile. For brevity, coefficients on the age vectors of the child and parents are suppressed in Table 2. The estimates based on father's earnings omit single-mother families, and vice versa, while the estimates including father's and mother's earnings separately include only two-parent families in which both parents work at some point. The last two estimates include all instances in which at least one parent works at some point (which encompasses more than 98 % of both sons and daughters with positive earnings) and combined earnings are the sum of both parents' earnings in any period in which at least one works.

In each variant in Table 2, the transmission elasticities from parental earnings or wages to those of their sons and daughters are precisely estimated but low. No matter which measure of parental earnings is adopted, transmission to the next generation remains small by comparison with estimates for other countries. In other words, these initial results are quite consistent with prior findings on Finland, summarized in Section 3.[22]

Table 3 shows OLS estimates of transmission from family income, with and without parental earnings included.[23] In all three specifications, the transmission elasticity from family income proves larger than the transmissions from parental earnings in Table 2. Moreover, the coefficient on family income proves relatively insensitive to the inclusion of parents' earnings in Table 3. For instance, in the last specifications in Table 3, the coefficient on family income exceeds that on earnings of the major earner with more than 99.9 % confidence for both sons and daughters. These estimates of the transmission elasticity from family income to earnings of the next generation are larger for

[21] Any distinction between this mean age measure, the mean age over which the individual is observed, or age at a specific point in time, is not always drawn in this context. However, in practice, the distinction has relatively little impact on the estimates.

[22] Projecting to age 40 and censoring earnings data, Jäntti et al. (2007) obtain slightly higher estimates for sons. The following section returns to the issue of an age interaction.

[23] Given the strong similarity in the results for earnings and wages in Table 2, for brevity, only results for the case of earnings are reported in the remainder of the paper. The specifications in family income include mean age of the father during periods when he reports positive income and a similar measure for the mother, both in polynomials through the fourth power, plus dummy variables for cases where no father or mother is initially present.

Table 3 Transmission from family income (between-individual estimates on positive earnings)

	Mean log earnings					
	Sons			Daughters		
Father's mean log earnings		0.012 (1.94)			0.002 (0.30)	
Mother's mean log earnings		0.006 (0.98)			0.015 (2.42)	
Major earner's mean log earnings			0.019 (2.62)			0.025 (3.19)
Family mean log income	0.225 (20.77)	0.234 (16.06)	0.212 (17.26)	0.173 (14.49)	0.174 (11.20)	0.153 (11.00)
No. of observations	14,364	11,346	14,139	13,860	10,993	13,645

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix

boys than for girls, but for both genders these transmission elasticities from family income are significantly larger than from parental earnings.[24]

Sensitivity analysis: alternative approaches to family income As anticipated in Eq. 20, the estimates of transmission from family income are sensitive to the number of observations on parents' incomes. This is explored more closely in the Electronic supplementary material—Appendix B, where it is demonstrated that the transmission elasticity for sons rises asymptotically and significantly with the number of observations on family income, with smaller increments the wider apart are these observations in time.

In Section 3, it was noted that incomes of both parents are aggregated in each time period even if they are not cohabiting at the time. It is shown in the Electronic supplementary material—Appendix B, however, that this is not a major concern; estimates of the transmission elasticity from family income prove quite insensitive to distinguishing states of cohabitation.

The distinction between parental earnings and family income is rendered possible in this study because of the importance of unearned income to a wide range of families. However, if this unearned income is driven by prior earnings, then the linear specification in Eq. 18 may not fully control for the effect of parental earnings (and hence inherited earning abilities) upon earnings of the next generation, which might act indirectly through the family income term.

[24] Several US studies also estimate transmission from family income to be larger than from parental earnings measures (see the review in Solon 1999). However, these studies uniformly concentrate upon transmission from parents' family income to child's family income, the latter presumably reflecting assortative mating in addition to the considerations in Section 1. As far as we are aware, no study incorporates both family income and parental earnings as regressors. A test for equality of all coefficients between the two stages of sampling within our data, based on the last specification in Table 3, gives $F(17, 14{,}105) = 0.937$ for sons and $F(17, 13{,}611) = 0.564$ for daughters. Pooling the subsamples appears not to be problematic.

The estimated weak associations with parental earnings even in the absence of family income measures, as reported in Table 2, suggest that this is not the case. However, it is possible to explore the issue further. Asset income may well reflect savings out of prior earnings, but it is state transfers rather than property income that are the major issue in this context; as noted previously, unearned income declines monotonically as a fraction of family income across almost the entire spectrum of income levels, with property income assuming importance only among the top two percentiles. However, receipts of state transfers in the forms of unemployment, maternity, and paternity benefits are also tied to prior earnings (see the notes on state transfers in the Electronic supplementary material—Appendix D). Before 1985, these transfers were quite tiny and none was taxable income and hence not part of our measure of family income. From 1987 onwards, our data report these transfers to each individual. Excluding these measures from family income diminishes the estimated transmission elasticity from family income, but only very slightly, by about 0.01 for both sons and daughters (see the results tabulated in the Electronic supplementary material—Appendix B). Retirement incomes are also tied to prior earnings, but separate data on these are not available. However, if the measure of family income includes only periods in which both parents earn (and hence are not retired), as well as subtracting the aforementioned unemployment and maternity benefits, the estimated transmission elasticities remain essentially unaffected.

Besides the method of averages, an alternative approach in estimating permanent family income is as a family fixed effect from panel estimation of log family income on the profile of ages reached by parents. The implication of adopting this alternative approach, within the first specification from Table 3, is actually to increase the estimated transmission elasticities from family income for both sons and daughters (see the results tabulated in the Electronic supplementary material—Appendix B).

Lastly, a number of studies have explored aspects of nonlinearity in intergenerational transmission elasticities (see, for instance, Corak and Heisz 1999; Grawe 2004; Bratsberg et al. 2007). The first linear specification from Table 3 is displayed in Fig. 1, along with an estimate that replaces the continuous measure of family mean log income with a vector of dummy variables for each percentile of family income. For neither gender is any clear departure from linearity apparent in Fig. 1.[25]

Sensitivity analysis: three generations Two issues are taken up here with respect to the influence of grandparents' incomes. The first arises with respect

[25] Restricting the specification to be linear, as opposed to the far more parsimonious form, generates an F statistics for sons of $F(97, 14{,}251) = 1.646$ and for daughters $F(97, 13{,}747) = 1.114$. A least absolute deviation estimate at the median also generates only a slightly lower value for transmission from family income than does the estimate at the mean. See Electronic supplementary material—Appendix B.

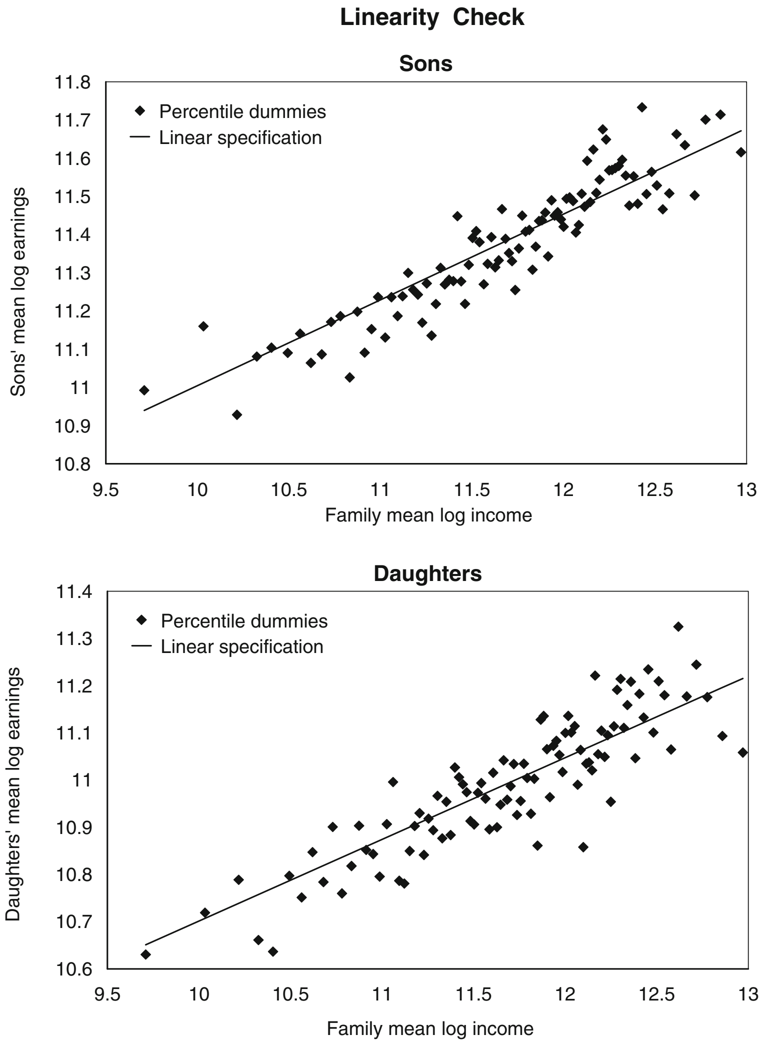

Fig. 1 Linearity check

to the endogenous taste hypothesis of Easterlin et al. (1980). Behrman and Taubman (1989) achieve a testable form of the fertility expression of this hypothesis; in the event that parents' taste for smaller families is positively shaped by higher incomes of their own parents, incomes of grandparents should be negatively associated with number of children in the third generation. To establish a parallel in the present context, suppose that the parental altruism parameter, α, depends upon grandparents' incomes. Since α is embodied in the intercept in the generic earnings equation for the child, at least in the credit-constrained context, incomes of grandparents should appear as an additional argument in shaping earnings of the third generation. In particular, if richer grandparents induce a taste for quality rather than quantity of children, the association with grandparents' incomes would be positive.

The second potential source of influence from grandparents' incomes arises from the use of Eq. 10 in deriving Eq. 13. This involves a tacit presumption that human capital investments in the parent, made by grandparents, are not credit-constrained. If, instead, the human capital of the parent were formed in a credit-constrained process, then using parental equivalents to the human capital choice (Eq. 6) and earning definition (Eq. 8) and inverting the latter yields

$$G_{\mathrm{p}} = E_{\mathrm{p}}\ Y_g^{-\rho^*} \left[\left(1+\rho^*\alpha^*\right) \left(\rho^*\alpha^*\right)^{-1} \right]^{\rho^*} \tag{21}$$

where Y_{g} represents the wealth of the grandparents. Using this in conjunction with Eqs. 6 and 8 then provides a second-order, auto-regressive, intergenerational transmission process:

$$E_{\mathrm{c}} = \kappa_2\ Y_{\mathrm{p}}^{\rho}\ E_{\mathrm{p}}^{\psi}\ Y_g^{-\psi\rho^*} \gamma_{\mathrm{c}} \tag{22}$$

where κ_2 is a function of α, ρ, α^*, ρ^*, and ψ. This formulation suggests that the logarithm of grandparents' income should have a negative influence upon log earnings of children in the third generation if ability transmission (ψ) is important. The intuition behind such a negative association derives from the proposition that richer grandparents would have invested more in the human capital of their children, given the abilities of these offspring to earn (G_{p}). Given E_{p}, G_{p} (and hence G_{c} and in turn E_{c}) would then be lower, the wealthier are the grandparents.[26]

Beyond interest in these two hypotheses themselves, an additional concern warrants noting. To the extent that parents' and grandparents' incomes are positively correlated, omission of the latter in examining intergenerational transmission would result in omitted variable bias, though of opposite direction depending upon which hypothesis dominates.

To explore this issue here, a separate data set on Finland is adopted. Statistics Finland manually matched a random 10 % sample of the individuals from the 1950 census with their later-established identity numbers. Combined with the same sampling frame of data described in Section 3, information on three generations of Finns is thus derived, with grandparents observed in 1950 and 1970 onwards, plus parents and their children observed from 1970 to 1999 (for greater detail on these data, see Pekkala and Lucas 2007).

Table 4 shows the results of OLS estimates on these data. An ambiguity arises with the measurement of Y_{g}, which may refer to income of the father's parents, the mothers' parents, or both. Accordingly, Table 4 includes each of these three variants. Given that only a 10 % sample was drawn from the

[26]Behrman and Taubman (1985) test a three-generational, schooling equivalent to Eq. 22 on a US sample. The same data set is used in Behrman and Taubman (1989) for a three-generational test of the Easterlin et al. (1980) fertility hypothesis. See also Robertson and Roy (1982), Warren and Hauser (1997), and Jeon and Shields (2005).

Table 4 Three generations

Sons' mean log earnings									
Family mean log income	0.249	0.254	0.210	0.243	0.244	0.206	0.296	0.290	0.209
	(27.91)	(27.46)	(12.43)	(29.82)	(28.84)	(15.11)	(11.02)	(9.94)	(4.01)
Father's parents' mean log income		−0.019	−0.019					−0.013	−0.012
		(2.68)	(2.69)					(0.60)	(0.55)
Mother's parents' mean log income					−0.006	−0.006		0.011	0.010
					(0.93)	(0.95)		(0.53)	(0.51)
Parent's mean log earnings			0.032			0.026			0.060
			(2.94)			(3.29)			(1.76)
No. of observations	27,289		27,281	31,853		31,730		2,905	
Daughters' mean log earnings									
Family mean log income	0.192	0.195	0.184	0.180	0.176	0.144	0.165	0.152	0.087
	(21.60)	(21.29)	(11.04)	(23.15)	(22.11)	(10.63)	(6.93)	(5.94)	(1.77)
Father's parents' mean log income		−0.011	−0.011					−0.013	−0.012
		(1.64)	(1.65)					(0.58)	(0.54)
Mother's parents' mean log income					0.008	0.008		0.058	0.056
					(1.22)	(1.20)		(2.74)	(2.69)
Parent's mean log earnings			0.008			0.025			0.049
			(0.77)			(3.14)			(1.59)
No. of observations	25,744		25,735	30,829		30,680		2,830	

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix

1950 census, the chances of identifying the parents of both father and mother, living in 1970, are low, and the resulting sample is therefore much smaller when incomes of both sets of grandparents are included. In each of the three variants, three specifications are reported. The first is a baseline, excluding incomes of grandparents, for comparison with Table 3. If anything, the estimated transmission from family income is slightly higher in the three-generation sample, though any differences are quite small. The second specification introduces the several alternative measures of grandparents' incomes. The only hint of a positive association is from mother's parents' incomes to earnings of daughters, when both sets of grandparents are included. Otherwise there seems little to suggest a positive association to support the Easterlin–Pollak–Wachter hypothesis. The third specification is a logarithmic transformation of Eq. 22, adding in parent's earnings (here interpreted as earnings of the major earner as in Table 3). Several of the coefficients on grandparents' incomes are negative, but only in the case of son's earnings is father's parents' income significantly below zero.

The three generation tests clearly do not prove particularly powerful in distinguishing the two hypotheses; indeed, it is feasible that the two simply offset each other. It is, however, important to note that inserting grandparents' incomes does little or nothing to diminish the estimated transmission from parental family income. Omitting grandparents' incomes does not appear to incur significant bias.

Sensitivity analysis: inclusion of zero earning observations A number of alternative representations of parental earnings have already been explored in connection with Table 2. However, Couch and Lillard (1998) raise an additional doubt about the validity of estimated intergenerational earnings elasticities that exclude observations when either the parent or child has zero earnings, an exclusion also imposed in our results so far. In the Electronic supplementary material—Appendix C, it is demonstrated that including zero earning observations in the Finnish context actually raises the intergenerational transmission estimates from family income but not from parental earnings. Although higher income families' children are more likely not to work while continuing their education, this effect is outweighed by the propensity of sons and daughters from lower income families to be unemployed or to drop out of the labor force (see also Ekhaugen 2009).

5 On the age dependence of intergenerational transmission

In the transition to an empirically tractable specification, it is assumed in Eq. 14 that the age profile of sons' and daughters' earnings is common to all individuals or at least that any variations are uncorrelated with parental earnings and incomes. Consider, however, the possibility of an interaction between age and the child's acquired human capital in shaping earnings.[27] Specifically, let the private returns on human capital asymptotically approach a limit as the person ages, such that

$$e_{ct} = \lambda_c + \lambda\ h_c \big/ a_{ct} + \mathbf{\Lambda}\ \mathbf{A}_{ct} + \xi_{ct} \tag{23}$$

where $h_c/\ell n H_c$. Given a common discount factor (δ_t) and life horizon among individuals, lifetime earnings may then be defined as the present value of e_{ct} such that

$$\ell n E_c = \Sigma\ \delta_t\ e_{ct} = \lambda_c + \lambda\Delta\ h_c + \lambda_0 \tag{24}$$

where $\Sigma\delta_t \equiv 1$, $\Delta = \Sigma\delta_t \big/ a_{ct}$, and λ_0 comprises the discounted elements of $\mathbf{A}_{ct}$ and ξ_{ct}. Substituting for λ_c from Eq. 24 in Eq. 23 yields

$$e_{ct} = -\lambda_0 + \ell n E_c - \lambda\Delta h_c + \lambda h_c / a_{ct} + \Lambda\ \mathbf{A}_{ct} + \xi_{ct}. \tag{25}$$

Substituting in Eq. 25 from Eqs. 11 and 12 in the case without credit constraints, or Eqs. 6 and 13 in the credit-constrained case, together with Eq. 15 then leaves an expanded version of the generic regression model (Eq. 18):

$$\begin{aligned} e_{ct} = {} & \beta_0 + \beta_1 e_{p\tau} + \beta_2 y_{p\tau} + \beta_3 e_{p\tau} \big/ a_{ct} + \beta_4 y_{p\tau} \big/ a_{ct} + \beta \big/ a_{ct} + \mathrm{B}\mathbf{A}_{ct} \\ & + B^* \mathbf{A}_{p\tau} + B^a \mathbf{A}_{p\tau} \big/ a_{ct} + \varepsilon_{ct}. \end{aligned} \tag{26}$$

In which,

Without credit constraint	With credit constraint
$\beta_1 = \psi\ (1-\rho*)\ (1-\rho)^{-1}\ (1-\lambda\Delta)$	$\beta_1 = \psi\ (1-\rho*)$
$\beta_2 = 0$	$\beta_2 = \rho - \lambda\Delta$
$\beta_3 = \psi\ (1-\rho*)\ (1-\rho)^{-1}\ \lambda$	$\beta_3 = 0$
$\beta_4 = 0$	$\beta_4 = \lambda.$

If $\lambda < 0$, then the intergenerational transmission, from parent's earnings in the wealth model and income in the credit-constrained case, should rise with age at which the child's earnings are observed.[28]

Age interaction estimates In a between-individuals estimator, any distinction between an age effect and a cohort (or time) effect cannot be discerned. Yet there are reasons to suspect that intergenerational economic mobility in Finland could differ across cohorts of children too, even within the time span

[27] See the discussion in Card (1999) for instance.

[28] Such an effect has previously been detected in the US context by Reville RT (Intertemporal and life cycle variation in measured intergenerational earnings mobility, unpublished), though Lee and Solon (2009) find less clear-cut patterns with respect to an age interaction. Equation 26 may be thought of as a specific form of the model developed in Haider and Solon (2006) which notes the potential age dependence of parameters such as β_1 and β_2 in Eq. 18. See also Jenkins (1987).

of our sample.[29] A panel estimation approach is therefore initially adopted to explore these potential age and cohort interaction effects.[30]

Only observations on positive earnings are included, in years in which the person is at least 21 years of age. The data refer to the continuous, annual portion of the sample from 1987 to 1999 and fixed effects are included for the year of observation. An error-components estimator is applied to these panel data in the first two results shown in Table 5, while the third adopts a maximum-likelihood estimator with an AR1 process in the errors. The estimated transmissions from the earnings of the father and mother in the first estimate, and of the major earner in the next two estimates, remain tiny and do not rise with age of the son or daughter. On the other hand, for both sons and daughters, the estimates indicate a significant and substantial rise in transmission elasticity from family income as the child ages. For both genders, similar results are found, though not shown, if a minimum age cutoff of 25 is adopted rather than 21.

To explore whether the interaction between family income and age reflects a cohort effect, rather than age, a term in family income relative to cohort is included in these panel estimates. Cohort is defined to equal 1 for sons and daughters who were 16 in 1970, through 17 for those aged less than 1 year in 1970. This cohort term proves to be tiny and virtually orthogonal to the other terms of interest.[31] The fixed effects for each year of observation (which are not tabulated) clearly reflect the substantial drop in real earnings during the massive recession after 1991. However, exclusion of these time fixed effects makes virtually no difference to the parameter estimates shown. Not surprisingly, given the orthogonality of the terms in family income relative to cohort and time fixed effects, the between-individual, OLS estimates in Table 5 also prove quite close to those obtained from the pooled, panel data.

Figure 2 illustrates the implied age profiles of the family income transmission elasticities from three of these estimates: random effects and OLS when earnings of both parents are included and the random effects case with earnings of the major earner (in which single-parent-earning families are included). Figure 2 also plots two additional error-components estimators, comparable to the first estimates in Table 5. In both additional plots, family

[29]For example, although all of the sons and daughters were potential beneficiaries of the guaranteed student loan scheme introduced in 1959 and explicitly subsidized after 1969, only the younger cohorts tended to benefit from the massive expansion in this system after the mid-1980s. Moreover, the various cohorts encountered the effects of the deep recession of the early 1990s at different ages (Hämäläinen and Hämäläinen 2005).

[30]Tests, using the *t*-bar statistics suggested by Im et al. (2003), reject with very high levels of confidence unit roots in the panel earnings data. See Electronic supplementary material—Appendix E.

[31]Pekkala and Lucas (2007) adopt more flexible forms of the cohort interaction and confirm the absence of any clear trend in family income transmission across cohorts born between 1954 and 1970 in Finland, though a downward decline is apparent among earlier cohorts born from 1930 to 1950.

Table 5 Age interaction effects

	Sons' log earnings					Daughters' log earnings				
	Panel			Time mean		Panel			Time mean	
	Random effects		AR1	OLS		Random effects		AR1	OLS	
Father's mean log earnings	0.017				0.014	0.008				0.006
	(4.02)				(3.04)	(1.89)				(1.28)
Mother's mean log earnings	0.008				0.004	0.015				0.016
	(1.91)				(0.96)	(3.64)				(3.44)
Major earner's mean log earnings		0.057	0.029	0.024			0.057	0.036	0.026	
		(4.05)	(28.77)	(3.94)			(3.62)	(31.90)	(4.15)	
Family mean log income	0.924	0.811	0.645	0.546	0.678	0.591	0.477	0.406	0.383	0.477
	(44.02)	(37.89)	(34.30)	(8.92)	(9.67)	(24.98)	(19.61)	(19.78)	(6.27)	(6.89)
Major earner's mean log earnings/child's age		−0.984	−0.178				−0.829	−0.202		
		(2.34)	(1.63)				(1.72)	(1.70)		
Family mean log income/child's age	−21.478	−18.419	−13.845	−10.691	−14.288	−13.591	−10.783	−9.086	−7.848	−10.183
	(37.36)	(29.43)	(23.97)	(5.74)	(6.74)	(20.55)	(15.02)	(14.46)	(4.32)	(4.97)
Family mean log income/cohort	−0.007	−0.004	−0.005			−0.005	−0.004	−0.004		
	(2.98)	(1.95)	(3.46)			(2.09)	(1.92)	(2.32)		
AR1 parameter			0.573					0.526		
			(409.5)					(346.4)		
No. of observations	299,637	349,022	349,022	31,182	26,644	280,082	327,165	327,165	29,905	25,556

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix

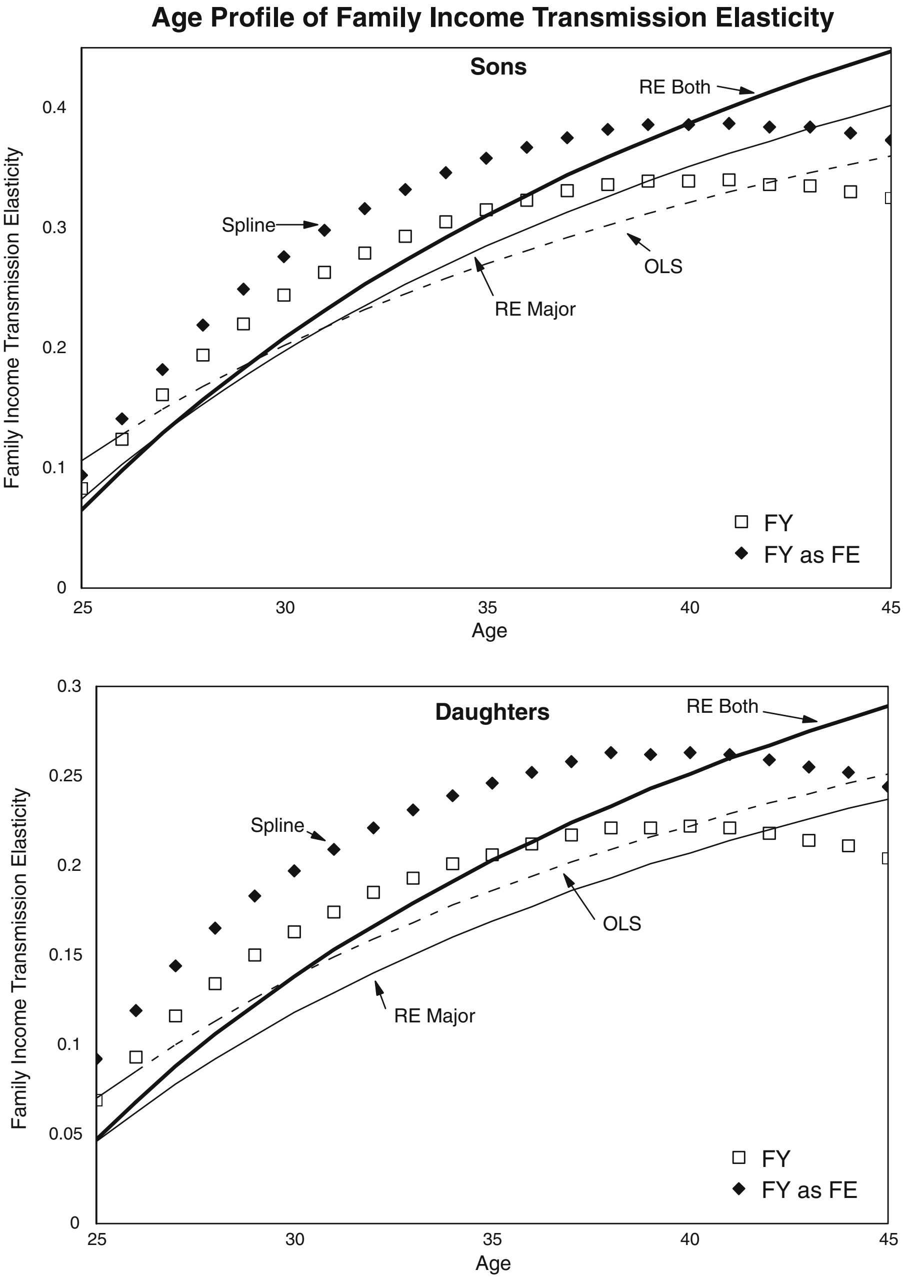

Fig. 2 Age profile of family income transmission elasticity

mean log income and family mean log income relative to child's age are replaced by family mean log income multiplied by a vector of dummies for each age at which the sons and daughters are observed. In one of these plots, the time mean of family log income is adopted, while in the other the estimate of family log income as a fixed effect, as described in the previous section, is deployed. These flexible formulations indicate a fairly monotonic rise in family income transmission elasticity, at least through about age 40, for both sons and daughters. Indeed, the hyperbolic specification adopted in Eq. 26 approximates this far more flexible age profile reasonably well.

Age and the returns to education It is hypothesized, in Eq. 25, that an interaction between human capital and age, in shaping earnings, underlies the rising transmission from family incomes as sons and daughters age. To explore this hypothesis a little more closely, let the logarithm of investment in the child's human capital (h_{c}) be approximated by the number of years of education completed. Interacting this measure of schooling with age at which earnings are observed, the first approach reported in Table 6 demonstrates that at least OLS estimates of returns to schooling indeed rise with age, for both sons and daughters in Finland.[32]

Such OLS estimates of the returns to education are, however, generally biased and inconsistent to the extent that schooling is correlated with omitted family background or ability effects. In an attempt to address this, a number of studies instrument years of schooling with some measure of educational reform or distance to school facilities (see Card 1999, for a critical review). The possibility of instrumenting years of schooling with institutional information about regional schooling in Finland was investigated. However, the available set of instruments proved too weak to offer meaningful results.[33] Instead, Table 6 reports OLS estimates on the first differences between the eldest and next oldest son and daughter pairs. To the extent that siblings share a common family environment and possibly similar abilities, any bias may be reduced in such comparisons. In fact, the results on first differences prove remarkably similar to the initial estimates. Moreover, if the set of observations is extended to include all sibling pairs of same gender, in descending age order, the results are hardly affected.

Both the influence of family income upon a child's earnings and the rate of return to schooling are thus estimated to rise with age of the child. To explore

[32]Controls are included, but not tabulated, for the terms a_{ct}^{-1}, $\mathbf{A}_{ct}$, $\mathbf{A}_{\mathrm{p}\tau}$, and $\mathbf{A}_{\mathrm{p}\tau}/a_{ct}$ from Eq. 26.

[33]The instruments available include the following: the proportions of school-age children in a region living 0–3, 3–5, or more than 5 km from a school in 1959 and in 1969; a dummy variable for whether the child was affected by the conversion to a comprehensive system, which was introduced on a rolling basis throughout the country, roughly from north to south, between 1972 and 1976 (cf. Black et al. 2005, on Norway; on the Finnish reform, see Pekkarinen et al. 2009); a dummy variable for whether the child's region of residence at age 18 contained a university town and, if not, distance to such a region (see Conneely and Uusitalo 1998). The partial F statistics for the seven available measures, in a first-stage regression including controls for a_{ct}^{-1} and $\mathbf{A}_{ct}$, are $F(7, 31{,}501) = 3.98$ for sons and $F(7, 30{,}227) = 2.81$ for daughters.

Table 6 Age interaction and the returns to education

	Sons' mean log earnings				Daughters' mean log earnings			
	OLS	Brothers difference	Recursive model	OLS	OLS	Sisters difference	Recursive model	OLS
Child's years of education	0.280	0.259		0.272	0.234	0.211		0.223
	(20.83)	(9.15)		(19.36)	(16.11)	(6.54)		(15.02)
Child's years of education/child's age	−6.867	−6.354		−6.884	−4.949	−4.365		−4.830
	(16.20)	(7.25)		(15.97)	(11.23)	(4.38)		(10.72)
Major earner's mean log earnings				0.027				0.030
				(4.50)				(4.87)
Family mean log income			0.562	0.243			0.367	0.143
			(9.32)	(3.85)			(6.10)	(2.32)
Family mean log income/child's age			−11.29	−3.209			−7.509	−2.675
			(6.13)	(1.66)			(4.19)	(1.45)
No. of observations	31,514	7,712	31,182	31,182	30,240	7,099	29,905	29,905

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix

further, the connection between these patterns considers the following simple, recursive model:

$$e_{ct} = \theta_0 + \theta_1 e_{p\tau} + \theta_2 s_c + \theta_3 s_c / a_{ct} + \theta_4 y_{p\tau} + \theta_5 y_{p\tau} / a_{ct} + \theta_6 / a_{ct} + \Theta \mathbf{A}_{ct} + \Theta^* \mathbf{A}_{p\tau} + \Theta^a \mathbf{A}_{p\tau} / a_{ct} + \zeta_{ct} \tag{27i}$$

$$s_c = \varphi_0 + \varphi_1 y_{p\tau} + \varphi_2 / a_{ct} + \Phi \mathbf{A}_{ct} + \Phi^* \mathbf{A}_{p\tau} + \Phi^a \mathbf{A}_{p\tau} / a_{ct} + \zeta_c \tag{27ii}$$

where s_c represents years of schooling for child c, ω_{ct} and π_c are stochastic disturbances, and other terms are previously defined. Why should Eq. 27i include terms in parental income ($y_{p\tau}$), given the child's schooling? At least two explanations have been offered in the literature. Mulligan (1999), upon finding that US children's wages are positively correlated with their parent's permanent income, given schooling of the child, interprets this in terms of a model with two dimensions of ability. The first dimension shifts the earnings function (as in our terms G_c and G_p); the other affects the returns to human capital (ρ^* for the parent). In the wealth model context, Mulligan then notes "Because {G_c} is positively correlated with {G_p and ρ^*}..., the earnings of adult children are positively related to parental income within a group of adult children whose parents made the same human capital investments."[34] Given that both G_p and ρ^* affect parental earnings (E_p), the additional dimension to ability then offers further justification for inclusion of $e_{p\tau}$ in Eq. 27i but not for inclusion of $y_{p\tau}$, given $e_{p\tau}$. Becker (1972), in commenting on related findings by Bowles (1972) for the USA, offers a second interpretation. Specifically, Becker suggests that the residual effect of parental income upon their child's earnings, controlling for the education level of the child, may simply reflect aspects of human capital not well represented by years of schooling. For instance, if the child's log human capital were defined by $h_c \equiv s_c + q_c$, where q_c might represent quality of schooling or postschool training, then the $y_{p\tau}$ terms in Eq. 27i would stem from the role of the omitted variable q_c in the credit-constrained case.

Estimates of Eq. 26, such as those reported in Table 6, may be thought of as reduced forms of Eq. 27 (Levine and Mazumder 2002) such that

$$\begin{aligned} \beta_2 &= (\theta_2 \varphi_1 + \theta_4) \\ \beta_4 &= (\theta_3 \varphi_1 + \theta_5)\,. \end{aligned} \tag{28}$$

After estimating both equations in Eq. 27 as seemingly unrelated regressions, the third columns in Table 6 report the estimates of the coefficients on $y_{p\tau}$ and $y_{p\tau}/a_{ct}$ implied by the formulae in Eq. 28. These derived estimates clearly resemble those obtained from direct estimation of the reduced form by OLS in Table 6. In fact, neither of the cross equation constraints in Eq. 28 can be rejected at a 99 % confidence level, either for sons or daughters.

[34]Mulligan (1999: S197). The portions in brackets are converted into notation of the present paper.

The final columns in Table 6 report between-individual, OLS estimates of Eq. 27i. The estimates of the returns to schooling and its age profile, θ_2 and θ_3, prove essentially orthogonal to inclusion of the additional terms in family income and major earner's earnings. The estimated coefficients on the major earner's mean log earnings remain tiny, suggesting that Mulligan's second dimension of ability may not be very important in this context. The estimated coefficients on family mean log income would, however, be consistent with Becker's suggested interpretation, in a credit-constrained case, that the quantity of schooling is an inadequate representation of human capital investments. More importantly, for present purposes, while the estimated returns to education continue to rise significantly with age, the estimated effect of family mean log income relative to child's age is statistically indistinguishable from zero, at least on a 99 % confidence test, both for sons and daughters. In other words, within this specification, after controlling for rising returns to schooling with age, no significant rise remains in the intergenerational transmission parameter as the child ages.

6 Within-generation credit constraints

The conventional focus on a budget constrained solely by permanent income implicitly assumes an absence of any credit constraint within the parental lifetime. For simplicity in reconsidering this presumption, distinguish two time periods: during and after the parents invest in the child's human capital. Let the parents' preferences be

$$U_{\mathrm{p}} = C_1 C_2^{\Gamma} Y_{\mathrm{c}}^{\alpha} \tag{29}$$

where C_i = parents' consumption in time period i. The budget constraints for the two time periods are

$$\begin{aligned} C_1 &\leq Y_1 - H_{\mathrm{c}} + L \\ C_2 &\leq Y_2 - L\,(1+\iota) \end{aligned} \tag{30}$$

where Y_i = parents' income in time period i, L are loans taken out by parents during period 1, and ι is the interest rate as before. If there is no constraint on L, then H_{c} depends only upon permanent income of the parents $\left(Y_{\mathrm{p}} = Y_1 + Y_2\,(1+\iota)^{-1}\right)$, as in Eq. 6. At the opposite extreme, if L is constrained to be zero, then Y_1 replaces Y_{p} in Eq. 6. With heterogeneous households, the mean family may exhibit a mix of these two extremes.

To explore this possibility empirically, a measure of the mean log family income during the interval when the child is age 0–17 is added to the generic regression model. In the first two OLS estimates in Table 7, the intergenerational transmission elasticity from mean log family income is found to be greater on family income received during this interval when the child is young,

Table 7 Family income during schooling years

	Sons' mean log earnings				Daughters' mean log earnings			
	OLS			IV	OLS			IV
Father's mean log earnings	0.008				0.003			
	(1.70)				(0.68)			
Mother's mean log earnings	0.003				0.016			
	(0.70)				(3.43)			
Major earner's mean log earnings		0.013	0.016	−0.013		0.022	0.022	−0.011
		(2.05)	(2.55)	(1.21)		(3.24)	(3.40)	(0.97)
Family mean log income	0.623	0.489	0.602	0.345	0.428	0.354	0.480	0.157
	(8.75)	(7.84)	(7.90)	(4.43)	(6.02)	(5.61)	(6.66)	(1.92)
Family mean log income/child's age	−13.939	−10.238	−13.994	−9.049	−9.490	−7.512	−11.661	−5.767
	(6.55)	(5.47)	(5.97)	(4.76)	(4.61)	(4.09)	(5.40)	(3.06)
Family mean log income when child 0–17	0.050	0.050	0.082	0.176	0.026	0.020	0.080	0.184
	(4.83)	(5.31)	(4.10)	(4.07)	(2.35)	(2.06)	(4.25)	(3.91)
Family mean log income when child 0–17/ no. of periods observed			−0.057				−0.092	
			(2.25)				(3.89)	
No. of observations	26,501	30,854	30,854	30,854	25,420	29,546	29,546	29,546
First-stage partial F				944.5				795.7

t statistics for a zero null hypothesis in parentheses. Standard errors from heteroskedastic-consistent matrix though these initial estimates are statistically weaker for daughters relative to sons. Family income is observed from one to five times before the children achieve age 18, depending upon the initial age of the child, with three or less observations for about 87 % of the sons and daughters. In the appendix, it is demonstrated that our prior estimates of intergenerational transmission elasticities increase significantly and asymptotically with the number of time observations over which family income is averaged, as suggested by Eq. 20. Accordingly, the third estimates in Table 7 append the time average of log family income when the child is less than 18 relative to the number of observations in forming this average (as well as a control for this number of observations). In doing so, the asymptotic transmission from additional income during schooling years of the child indeed proves larger for both genders.

Parents who are anxious to invest more in their child's human capital may attempt to load their income into the earlier part of the child's life. Accordingly, in the final results in Table 7, mean family income when the child is age 0–17 is instrumented, adopting a variable based on the average unemployment rates of fathers and mothers living in the same region as the parents of the observed son or daughter. For whichever parent is the major earner, the gender-specific, regional rate is averaged over periods in which the particular parent contributes to family income. The instrument is then the difference in this mean unemployment rate during the period when the child is age 0–17 as opposed to the average over the entire period of observation (Erola and Moisio 2002). As shown in Table 7, this instrument proves very powerful in the first stage. To the extent that the instrumental variable is valid, the results suggest that ordinary least squares underestimates the addition to intergenerational transmission from family income during schooling years.

In a US context, Haider and Solon (2006) and Grawe (2006) note that the attenuation inconsistency, from estimating intergenerational earnings transmission by ordinary least squares on fathers' current earnings as a proxy for lifetime earnings, is minimized if fathers are observed in the age range from their early 30s to mid-40s. Böhlmark and Lindquist (2006) find comparable results for Sweden (see also Baker and Solon 2003, on Canadian men). Since the average age of our Finnish major earners is 42 when their children are of school age, the larger transmission elasticity from family income during these schooling years may simply reflect an effect of diminished attenuation inconsistency during this period of observation, rather than intertemporal borrowing constraints. Towards exploring this distinction, the prior specifications may therefore be extended to include an interaction between family mean log income when the child is ages 0–17 and the square of the difference from 40 in the major earner's mean age during these years (plus a control for this age differential term itself). In both the ordinary least squares and instrumental variables estimates, the estimated coefficients on the additional interaction term are indistinguishable from zero, even at a 90 % confidence level. In other words, this supports the notion that additional family income when both sons and daughters are passing through school is indeed associated with higher subsequent earnings for this next generation.

7 In perspective: a summing up

A nested model has been developed, incorporating an auto-regressive, intergenerational process in earnings that stems from heritability in earning abilities, and a budget constraint on human capital investments in children imposed by the permanent incomes of their parents. Any transmission from parental earnings to those of their children is found to be very weak in the rich Finnish data, while the transmission from family incomes to earnings of both sons and daughters is significantly and substantially larger. Moreover, the income available to parents during the phase when their children are passing through school is found to have an additional effect on the future earnings of both sons and daughters, even after instrumenting this additional income. The influence of permanent income is consistent with constraints on the capacity of parents to borrow against future incomes of their children; the additional influence of family income during the child's early life is consistent with constraints on parents' ability to raise credit against their own subsequent incomes.

Exploration of a three-generation data set points to three things: a lack of support for the Easterlin–Pollak–Wachter hypothesis of endogenous tastes, which might have offered an alternative explanation for the transmission from family income to children's earnings; further support for a weak ability transmission parameter; and confidence that omission of grandparents' incomes does not substantially bias estimates of transmission from parents' incomes.

The lack of any substantial, auto-regressive process from parent's earnings to those of their offspring in Finland proves insensitive to whether both father's and mother's earnings are considered in two-earner families, or the earnings of the higher earning parent are adopted across both single-parent and two-parent families, and insensitive to several other combinations of these parents' earnings too. The significant intergenerational transmission from parents' family income to earnings of their sons and daughters also proves robust to a range of permutations considered: the estimated elasticity of transmission increases significantly the more observations are available on family income, though it proves even greater if permanent income is estimated as a fixed effect instead; the elasticity is fairly insensitive to alternative definitions of family income when parents separate and to exclusion of earnings-related maternity and unemployment benefits and of retired persons; moreover, a less parsimonious specification suggests that the elasticity does not vary greatly across the income spectrum. In addition, the elasticity is estimated to increase if zero-earning observations of sons and daughters are incorporated because, although higher income families' children are more likely not to work while continuing their education, this effect is outweighed by the propensity of sons and daughters from lower income families to be unemployed or to drop out of the labor force.

Sensitivity of intergenerational mobility estimates to age of the second generation has been noted in prior work on other countries. In the Finnish data, at least through about age 40, a fairly monotonic rise in the intergenerational transmission elasticity from family incomes is observed, the older the age at which both sons' and daughters' earnings are observed. Moreover, both in cross section of sons and daughters and by comparisons of siblings, the returns to education are shown to rise with age in Finland. After decomposing the intergenerational transmission elasticity to account for this rise in the rate of return to years of schooling, as sons and daughters mature, any residual rise in transmission elasticity with respect to age proves statistically insignificant. Having wealthy parents increasingly pays off as one becomes older, and this is substantially because the returns to parental investments in years of schooling rise with age.

Because the income distribution is fairly even in Finland, there are relatively few extremely wealthy families. This may diminish the likelihood of observing families able to leave positive bequests, as required for the wealth model to hold. Given the prevalence of collective bargaining in Finland, earning compression, particularly in the earlier years of our data, may render earnings less likely to reflect idiosyncratic abilities (Eriksson and Jäntti 1997). Nonetheless, the evidence on the importance of family incomes is compelling. In other countries, unearned income might reflect abilities to invest, which could be passed on as an ability to earn in the next generation. In Finland, however, most of unearned income is derived from state transfers rather than investment income. Yet, even in the presence of Finland's comprehensive welfare state system, families' permanent incomes, as well as their incomes during children's schooling years, appear important in understanding the substantial intergenerational income intransigence.

Finally, as noted in the introduction, prior work on the Nordic countries (including Finland) suggests low transmission rates from parent's earnings to those of their children. The results in this paper certainly concur with these impressions; intergenerational auto-regression in earnings is low in Finland, but transmission from family income to children's earnings is not. International comparisons are difficult. Nonetheless, our estimates of intergenerational transmission in Finland are not obviously lower than those obtained outside of the Nordic region. The privileges that come with higher incomes, not only schooling but the many constituents of human capital, appear to be important factors shaping the earnings of the next generation in Finland. In contrast, all of our estimates point to small and fairly insignificant values for a parameter of inherited earning ability. Apples indeed fall close to Finnish trees; having either a silver or golden spoon handy is what breeds success.

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